Land Retirement Program Design in the Presence of Crop Insurance Subsidies

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Abstract

The U.S. federal government implements environmental, biofuels and crop insurance programs that influence land use. They are not well integrated in that cost savings from crop insurance subsidies are not acknowledged when screening land for retirement or when calculating the cost of land retirement programs. We identify and evaluate an optimal benefit index for enrollment in a land retirement program that includes a sub-index to rank land according to insurance subsidy savings. All else equal, land ranked higher in the Lorenz stochastic order should be retired first.

Keywords: agro-environmental policy, budget, Conservation Reserve Program, crop failure, Environmental Benefit Index, Lorenz order.

JEL classification: Q18, Q28

Introduction

The United States federal government influences land use in many ways. Perhaps most directly, it supports the Conservation Reserve Program (CRP) and other programs to remove land from agricultural production and directly influence environmental benefits from farmed land. The 2007 farm bill allows for about $5 billion per year in such expenditures over the period 2008-17. Through biofuels use mandates and other means, the federal government has increased demand for cropland. In addition, under the Agricultural Risk Protection Act (ARPA) of 2000 it provides subsidies on crop insurance. Over the years 2000-07, U.S. taxpayers have transferred about $11 billion to farmers in this way (Babcock 2008). With multiple goals that include transfers, land use policy design is inevitably a challenging endeavor where one difficulty is to integrate into a coherent framework disparate policy interventions that affect land use.

Consider how land retirement and crop insurance programs integrate. The CRP is targeted at removing cropland from production if it performs sufficiently well on the program's Environmental Benefits Index (EBI). Factors entering the EBI include benefits to wildlife, water quality metrics, air quality consequences, erosion propensity and carbon sequestration potential (Hajkowicz, Collins and Cattaneo 2009). Enrollment cost factors are also included where, *ceteris paribus*, land that commands higher market rent will perform worse on the index and so is less likely to be accepted for enrollment.

Omitted from the index, however, is the reduction in premium subsidies from crop insurance that would occur were the land to be removed from production. This is an important omission, as a cursory inspection of any program enrollment map shows high concentration of enrolled acres in the southern Corn Belt, the eastern Dakotas, Montana, the southern Great Plains and parts of the Palouse (Ribaudo *et al.* 2001). These are in the main marginal cropland regions where CRP enrollment costs are low and benefits may be high. For example, these

lands tend to be erosion-prone and more vulnerable to nutrient losses that ultimately pollute water bodies.

What makes these locations marginal for cropping and environmentally sensitive often also makes them poor crop insurance prospects (Lubowski *et al.* 2006). As one instance, low organic matter content renders soil susceptible to wind and water erosion. In addition, low organic matter soil is poorly aerated and incapable of retaining rainfall to buffer crops against drought (p. 70 in Smil 2001). Thus, low mean yield land typically also displays comparatively high yield variance as reflected in coefficients of variation (p. 103 in Woodard 2008). If crop insurance receives a percentage of premium subsidy as under ARPA, and even if it receives a fixed per acre subsidy as was the case before ARPA, then marginal but uncropped acres are likely to be drawn into production. In addition to placing constraints on attaining environmental goals, the policy will involve a cost to the budget.

It should be no surprise then that the ratio of indemnities paid out over farmer premiums paid in during 2000-07 was above two in Oklahoma, Montana, Texas, Kansas and the Dakotas (Babcock 2008). The ratio was less than one in the prime cropland states of Iowa, Illinois and Indiana. Transfers are large, where total indemnities less farmer premiums paid over the period 2000-07 exceeded $1 billion each for Texas, Kansas, North Dakota, and South Dakota.[1] Consistent with findings in the small body of work on the issue, Lubowski *et al.* (2006) identified a modest but definite impact of crop insurance subsidies in promoting cropping on environmentally sensitive land.

Our concern is with the omission of crop insurance subsidies as an avoided cost when constructing an index to use for enrollment. There are two reasons this should be a concern. Although program fund sources differ, federal taxes spent and saved have equal weight when calculating the budget deficit. Secondly, were avoided crop insurance costs to be included in

[1] Glauber (2004), over 1981-2003, and Woodard (2008), over 1980-2006, provide similar insurance loss data for the Corn Belt states.

the index, then incentives for optimal land allocation would be strengthened. Inclusion would mitigate dissonance across the suite of agro-environmental policies.

This paper does four things. It identifies precisely how crop insurance savings should be included in a modified EBI. It studies the nature of the ordering that should be used to rank crop insurance savings from land retirement. It evaluates how the index should be weighted absent a metric for these savings. And it discusses why the EBI index omits avoided subsidy costs.

Model

The model used in this work is an extension of the standard constrained benefit maximization specification, as used in, e.g., Babcock *et al.* (1996), Newburn, Berck, and Merenlender (2006), and Feng *et al.* (2006). Define by $k \in \{1, 2, \ldots, K\} \equiv \Omega$ the *k*th acre of land, all of unit size which we set as one acre. The set of all subsets of Ω is written as $\mathcal{P}(\Omega)$. Environmental benefits arising from retirement amount to e_k while production costs avoided by placing the land in retirement equal c_k. We write net benefits as $b_k = e_k - c_k$. Therefore if set $\mathcal{S} \subseteq \mathcal{P}(\Omega)$ is placed in retirement then net benefits amount to $\sum_{k \in \mathcal{S}} b_k$.

The land management planner faces a constraint and a cost when seeking to maximize the value of $\sum_{k \in \mathcal{S}} b_k$. The constraint is the requirement that expected total forgone agricultural production equals level Q where stochastic yield on the *k*th acre is q_k and mean yield is μ_k. Formally, $Q = \sum_{k \in \mathcal{S}} \mu_k$. At the practical level, this constraint reflects pressure from animal agriculture, the biofuels industry and ultimately commodity end consumers regarding acres available for production. The cost is the social welfare deadweight loss from raising the funds used to retire the land. This has two components, where the first is rent r_k that would have to be paid to temporarily retire the land. But there are also savings that lower the level of funds

required. These savings arise from avoiding the cost of subsidizing crop insurance.

We model the insurance indemnity as follows. The *k*th acre is insured for losses below yield $\beta\mu_k$ where $\beta \geq 0$ and the price paid is p. With expectation operator $\mathbb{E}[\cdot]$ over the *k*th acre yield distribution, the expected indemnity is $p\mathbb{E}[\max(\beta\mu_k - y_k, 0)]$. Given subsidy rate s, the expected cost to the government is $sp\mathbb{E}[\max(\beta\mu_k - y_k, 0)]$. The cost would be avoided were the land taken out of production. To be succinct we write $\delta_k(\beta) = \mathbb{E}[\max(\beta\mu_k - y_k, 0)]$, or just δ_k when β is understood to be fixed.

Both rent r_k and government indemnity payments are transfers. If surpluses to all economic agents are weighted equally then the social cost of making these payments equals the deadweight loss associated with raising taxes. As is standard in the literature, write this as a price τ so the net cost is

(1) $$\tau \times \left(\sum_{k \in \mathcal{S}} r_k - sp \sum_{k \in \mathcal{S}} \delta_k(\beta) \right).$$

The land planner's problem is[2]

(2) $$\max_{\mathcal{S} \in \mathcal{P}(\Omega)} \quad \sum_{k \in \mathcal{S}} b_k + \lambda \times \left(Q - \sum_{k \in \mathcal{S}} \mu_k \right) - \tau \times \left(\sum_{k \in \mathcal{S}} r_k - sp \sum_{k \in \mathcal{S}} \delta_k(\beta) \right),$$

where λ is a Lagrange multiplier representing the shadow value of output. From (2) and writing

(3) $$U_k(\beta) = \hat{b}_k - \tau \hat{r}_k + \tau sp \hat{\delta}_k(\beta); \qquad (\hat{b}_k, \hat{r}_k, \hat{\delta}_k(\beta)) = \left(\frac{b_k}{\mu_k}, \frac{r_k}{\mu_k}, \frac{\delta_k(\beta)}{\mu_k} \right);$$

the inclusion criterion for the *k*th acre is

[2] The standard approach is to maximize over a closed bounded parcel size interval, say, $[0, \overline{x}_k]$, so that the problem is a mixed (discrete and continuous) linear program. Except for the last acre, interior solutions will not occur, so we have dropped the cumbersome notation associated with partial allocation of an acre and focus on solving a discrete linear program.

$$(4) \qquad \text{include} \quad \begin{cases} \text{all acre if} & U_k(\beta) > \lambda; \\ \text{none of acre if} & U_k(\beta) < \lambda; \\ \text{some of acre if} & U_k(\beta) = \lambda. \end{cases}$$

We see then that three factors enter into the determination of $\mathcal{S} \subseteq \mathcal{P}(\Omega)$. Acres having larger environmental benefits per unit of expected forgone production, i.e., larger $\widehat{b}_k$, are more likely to merit inclusion in the retirement program. Acres with larger rental cost per unit expected forgone production, i.e., larger $\widehat{r}_k$, are less likely to merit inclusion. If the social cost of raising taxes, τ, is high then $\widehat{r}_k$ will weigh heavily when compared with an acre's $\widehat{b}_k$ value in the inclusion decision. Finally acres with higher expected subsidy costs per unit expected forgone production, i.e., higher $\widehat{\delta}_k(\beta)$, should be included. Here the weighting relative to $\widehat{b}_k$ is τsp and not just τ so that percent subsidy and price that crop losses receive also matter.

Insurance Loss Index

Defining $\widehat{\delta}_k(\beta)$ as the Insurance Loss Index (ILI), we may write

$$(5) \qquad \widehat{\delta}_k(\beta) = \frac{\mathbb{E}[\max(\beta\mu_k - y_k, 0)]}{\mu_k}.$$

Several comments are in order concerning how yield affects this measure and, so, index U_k. First, suppose that the *j*th and *k*th acres are stochastically ordered such that the *j*th is larger in the first-order sense.[3] Then $\mu_j \geq \mu_k$. But we cannot establish *a priori* whether

$\mathbb{E}[\max(\beta\mu_j - y_j, 0)] \leq \mathbb{E}[\max(\beta\mu_k - y_k, 0)]$ or $\widehat{\delta}_j(\beta) \leq \widehat{\delta}_k(\beta)$.

To explore the issue further, consider a two-point yield distribution for the *j*th acre. The

[3] Distribution $F_j(\cdot)$ is larger than distribution $F_k(\cdot)$ in the first-order sense whenever $\mathbb{E}[g(y_j)] \geq \mathbb{E}[g(y_k)]$ for all increasing functions $g(\cdot)$. Alternatively, $F_j(y) \leq F_k(y)$ on the union of supports for y. Stochastic dominance orderings are explained in chap. 3 of Gollier (2001).

distribution has yield y_j^l with probability $\pi \in (0,1)$ and yield y_j^h with probability $1-\pi$ where $y_j^h > y_j^l$. The *k*th acre distribution has yield $y_k^l = y_j^l$ with probability π and yield $y_k^h = y_j^h + \phi$ with probability $1-\pi$ where $\phi > 0$. We refer to the distribution shift from *j*th to *k*th acre as an *upper stretch* because the high yield is shifted whereas the low yield remains fixed. It is a restricted version of first-order stochastic dominance.[4]

Result 1: If the *k*th acre first-order dominates the *j*th acre by an upper stretch then $\hat{\delta}_k(\beta) \geq \hat{\delta}_j(\beta)\ \forall \beta \in [0,1]$.

Requirement $\beta \leq 1$ is not onerous in that $\beta > 1$ would involve a yield guarantee larger than mean yield. The result has yield becoming more variable while mean yield increases, where the former effect dominates to drive up the ILI.

Alternatively, consider the same two-point yield distribution for the *j*th acre but let the *k*th acre distribution have yield $y_k^l = y_j^l + \phi$ with probability π and yield $y_k^h = y_j^h$ with probability $1-\pi$ where $\phi \in (0, y_j^h - y_j^l)$. We refer to the distribution shift from *j*th to *k*th acre as a *lower contraction*, and it is also a restricted version of first-order dominance.

Result 2: If the *k*th acre first-order dominates the *j*th acre by a *lower contraction* then $\hat{\delta}_k(\beta) \leq \hat{\delta}_j(\beta)\ \forall \beta \in [0,1]$.

The result has yield becoming less variable while mean yield increases, and in light of (5) it may not be a surprise to learn that the dominating shift drives down the ILI. The pair of results clarify that, *ceteris paribus*, ranking acres for land retirement using the first-order dominance criterion may not be efficient.

[4] Where not obvious, results are demonstrated in the appendix.

Suppose instead that the *j*th and *k*th acres are stochastically ordered such that the *k*th's yield distribution is a mean-preserving contraction relative to that for the *j*th acre.[5]

Result 3: If the *k*th acre dominates the *j*th acre by a mean-preserving spread then $\hat{\delta}_k(\beta) \geq \hat{\delta}_j(\beta)\ \forall \beta \in [0,1]$.

Upon controlling for the mean, the more risky distribution should be ranked higher for inclusion in the land retirement program.

We also ask what effect an independent risk source would have on whether to retire an acre. For example suppose two acres are equivalent in all ways except that one is on a flood plain and the other is not. Let $y_k \overset{\text{dist}}{=} By_j$ where B, independent of y_j, is Bernoulli having value 1 with probability π and value 0 (i.e., the flood event) otherwise. The distributions of y_j and y_k are not ordered by a mean-preserving spread as the mean is not preserved. The distributions are not ordered by the upper stretch or lower contraction conditions either as the distribution of y_j is arbitrary and not two-point. We have the following:

Result 4: Under the above representation of disaster risk, $\hat{\delta}_k(\beta) \geq \hat{\delta}_j(\beta)\ \forall \beta \geq 0$.

Being exposed to that independent risk, the *k*th acre of land is indeed riskier and, *ceteris paribus*, should be included in the retirement program whenever the *j*th acre is.

[5] Distribution $F_j(\cdot)$ is a mean-preserving contraction (mpc) when compared with distribution $F_k(\cdot)$ whenever $\mathbb{E}[g(y_j)] \leq \mathbb{E}[g(y_k)]$ for all convex functions $g(\cdot)$. The reverse of an mpc is a mean-preserving spread (mps), meaning $F_j(\cdot)$ is an mps when compared with $F_k(\cdot)$ whenever $\mathbb{E}[g(y_j)] \geq \mathbb{E}[g(y_k)]$ for all convex functions $g(\cdot)$.

Lorenz Order

While appealing to standard stochastic dominance tools in the economics of uncertainty, the comments made to this point concerning how an acre's yield risk profile affects the ILI are not precise in that results 1-4 do not cover all cases. The conditions for the ILI to increase are sufficient but they are not necessary. Linear homogeneity of the function $\max(x,0)$ allows us to identify the precise stochastic order that ranks acres across all values of $\beta \geq 0$. From (5) we may write

(6) $$\hat{\delta}_k(\beta) = \mathbb{E}[\max(\beta - z_k, 0)]; \qquad z_k = \frac{y_k}{\mu_k}.$$

The Lorenz stochastic order provides much insight on this relation. Shaked and Shanthikumar (2007, p. 116) report:

Definition 1: For non-negative random variables y_j and y_k, y_k is said to dominate y_j in the Lorenz order (written as $y_k \ {}_{\text{Lor}}\geq y_j$) whenever the mean-normalized random variable y_k / μ_k is a mean-preserving spread of y_j / μ_j.

The stochastic ordering is equivalent to the Lorenz order so familiar to students of income distribution orderings (Atkinson 1970). We are considering yields, non-negative random variables. For non-negative random variables, Arnold (1986, p. 37) reports that $y_k \ {}_{\text{Lor}}\geq y_j$ and $\hat{\delta}_k(\beta) \geq \hat{\delta}_j(\beta)\ \forall \beta \geq 0$ imply each other. Equivalently, $y_k \ {}_{\text{Lor}}\geq y_j$ whenever $\mathbb{E}[g(y_k / \mu_k)] \geq \mathbb{E}[g(y_j / \mu_j)]$ for all continuous convex functions $g(\cdot)$. Therefore, relation (6) allows us to state:

Result 5: If $y_k \ {}_{\text{Lor}}\geq y_j$ then $\hat{\delta}_k(\beta) \geq \hat{\delta}_j(\beta)\ \forall \beta \geq 0$.

Lorenz curves provide equivalent information for mass distributions, not probability

distributions. These curves have been applied elsewhere as a tool for analyzing and explaining the effects of environmental programs. Babcock *et al.* (1996, 1997) use them to compare budget effectiveness across disparate CRP enrollment criteria. Groot (2009) uses them to compare country population shares with global greenhouse gas emission shares.

Some examples can illustrate how the order ranks yield distributions. If $y_k \geq 0$ and $y_j \overset{\text{dist}}{=} y_k + \theta$, $\theta > 0$, then $y_k \underset{\text{Lor}}{\geq} y_j$ and the *k*th acre scores higher on the index for acres to be retired.[6] This makes sense as the *k*th acre more readily satisfies the commodity supply condition while being possessed of the same variability. Alternatively, if $y_j \overset{\text{dist}}{=} \theta y_k$, $\theta > 0$, then $\hat{\delta}_j(\beta) \equiv \hat{\delta}_k(\beta)\ \forall \beta \geq 0$ and the distributions are equal under the ordering. We will have reason to comment on this observation at a later point.

More generally, Wilfling (1996) has considered the pair of positive, continuous, increasing functions $g(r)$ and $h(r)$ on $r \geq 0$ with $g(r) > 0$ and $h(r) > 0$ on $r > 0$. If $h(r)/g(r)$ is increasing then $h(r) \underset{\text{Lor}}{\geq} g(r)$. For example, let $y_j = g(r) = \theta_g r^\alpha + \tau_g$ and $y_k = h(r) = \theta_h r^\alpha + \tau_h$ where parameter set $\{\theta_g, \theta_h, \tau_g, \tau_h, \alpha\}$ is positive and $r \geq 0$ is growing season rainfall on arid land. Then the derivative of $h(r)/g(r) = [\theta_h r^\alpha + \tau_h]/[\theta_g r^\alpha + \tau_g]$ has the sign of $\theta_h/\tau_h - \theta_g/\tau_g$. So if $\theta_h/\tau_h > \theta_g/\tau_g$ then $y_j \underset{\text{Lor}}{\geq} y_k$ and the *j*th acre is the better candidate for including in the program.

One may also apply Wilfling's observation to the role of inputs on the ILI. Writing x as an input, let output be $f(r,x)$ with r random, and we study $f(r, x = x_2)/f(r, x = x_1)$ where $x_2 > x_1$. The yield distribution under x_2 is larger (smaller) in the Lorenz order and so more (less) costly to subsidize per bushel forgone, if

[6] Here $x_1 \overset{\text{dist}}{=} x_2$ means that the distributions of x_1 and x_2 are equal almost everywhere.

$\partial \ln[f(r, x = x_2)] / \partial r > (<) \partial \ln[f(r, x = x_1)] / \partial r$. When x is chosen over an interval, say, $[\underline{x}, \overline{x}]$, this requires that $\partial^2 \ln[f(r, x)] / \partial r \partial x > (<) 0$ over the interval, i.e., that the production function be log-supermodular (log-submodular) (Athey 2002). Intuitively, a higher value of x makes the land more rainfall sensitive. The yield distribution is more variable in proportional terms, where proportionality normalizes away the effects of a shift in the mean yield, as does the Lorenz ordering.

Relation to Just-Pope Technology

In a widely applied innovation, Just and Pope (1979) endogenized yield riskiness by representing input-conditioned yield as $y(x) = g(x) + h(x)\varepsilon$. Here $\mathbb{E}[\varepsilon] \equiv 0$, ε follows some distribution function $F(\varepsilon)$, x is a vector (possibly scalar) of inputs while $g(x)$ and $h(x)$ are assumed to be positive. If $h(x)$ is increasing in an input then that input is held to be risk increasing.

Since $\mathbb{E}[y(x)] = g(x)$, it follows that the input-conditioned ILI is

$$(7) \qquad \hat{\delta}(\beta; x) = \frac{\mathbb{E}[\max(\beta g(x) - y(x), 0)]}{g(x)} = \mathbb{E}[\max(\psi - m(x)\varepsilon, 0)],$$

where $\psi = \beta - 1$ and $m(x) = h(x) / g(x) > 0$. If $m(x)$ is increasing then an increase in the level of inputs induces a mean-preserving spread in random variable $m(x)\varepsilon$ and so increases the value of $\hat{\delta}(\beta; x)$. More formally, we can write the result as follows.

Result 6: Consider the Just-Pope technology as presented above. If $x_b \geq x_a$ then $m(x)$ increasing implies $\hat{\delta}(\beta; x_b) \geq \hat{\delta}(\beta; x_a)$ while $m(x)$ decreasing implies $\hat{\delta}(\beta; x_b) \leq \hat{\delta}(\beta; x_a)$.

An increasing $m(x)$ function would suggest that a restriction on inputs, perhaps due to environmental damage considerations, would decrease the ILI. The restriction would make the

acre less costly in terms of insurance subsidy costs per expected bushel forgone. But of course the expected cost of attaining a given expected yield would increase.[7]

If we suppose that the Just-Pope technology is appropriate, then do inputs have a uniformly positive or negative $m(x)$ function? Often evidence on the $m(x)$ function as reported in the literature is insufficient to inform on the function's monotonicity status. A set of functional forms that allow for this arises when both $g(x)$ and $h(x)$ are Cobb-Douglas, for then the ratio is also Cobb-Douglas and monotone without regard to the relevant input domain.

Using Day's (1965) classic data set on yield and nitrogen level, Just and Pope (1979) do estimate this technology for corn and oats experimental plot data. For corn the estimation is $y = Ax^{0.353} + Bx^{0.127}\varepsilon$ where $A > 0$ and $B > 0$ are coefficients and x represents nitrogen. Thus, nitrogen is risk increasing in the Just-Pope sense. Nonetheless, $m(x) = x^{0.127} / x^{0.353} = x^{-0.226}$, a decreasing function, so that the estimates suggest an increase in nitrogen *reduces* the ILI. For oats, $y = Ax^{0.310} + Bx^{0.200}\varepsilon$ so that nitrogen also reduces the ILI even though it would be viewed as a risk increasing input.

Consequences of Excluding ILI from EBI

Consider now the case when direct information on $\delta_k(\beta)$ is unavailable. Assuming a continuous distribution on acre attributes, and so dropping subscript *k*, we suppose that variable set $(\widehat{b}, \widehat{r}, \widehat{\delta})$ is joint normal with means $(\varsigma_b, \varsigma_r, \varsigma_\delta)$, deviations $(\eta_b, \eta_r, \eta_\delta) = (\widehat{b} - \varsigma_b, \widehat{r} - \varsigma_r, \widehat{\delta} - \varsigma_\delta)$ and covariances

[7] To also accommodate social planner choice over inputs, problem (2) should be extended. The maximization should be over $\mathcal{S}$ and x where b_k, μ_k, and $\delta_k(\beta)$ depend on x. Benefits on farmed acres should also be included, as should the cost of x.

(8) $$\begin{pmatrix} \sigma_{bb} & \sigma_{br} & \sigma_{b\delta} \\ \sigma_{br} & \sigma_{rr} & \sigma_{r\delta} \\ \sigma_{b\delta} & \sigma_{r\delta} & \sigma_{\delta\delta} \end{pmatrix}.$$

Letting $\chi = \sigma_{bb}\sigma_{rr} - \sigma_{br}^2 > 0$, then the conditional mean of $\hat{\delta}$ is given as (p. 34 in Tong 1990)

(9) $$\mathbb{E}[\delta \mid \hat{b}, \hat{r}] = \varsigma_\delta + \rho_{b\delta|r}\eta_b + \rho_{r\delta|b}\eta_r; \quad \rho_{b\delta|r} \equiv \frac{\sigma_{b\delta}\sigma_{rr} - \sigma_{br}\sigma_{r\delta}}{\chi}; \quad \rho_{r\delta|b} \equiv \frac{\sigma_{r\delta}\sigma_{bb} - \sigma_{br}\sigma_{b\delta}}{\chi}.$$

Here, $\rho_{b\delta|r}$ and $\rho_{r\delta|b}$ are, respectively, the *r*- and *b*-conditioned partial correlations between δ and the other observed variable. Insert into expectation of index function (3) conditional on knowing $\hat{b}$ and $\hat{r}$ to obtain

(10) $$\begin{aligned} &\mathbb{E}[U \mid \hat{b}, \hat{r}] = \hat{b} - \tau\hat{r} + \tau sp\mathbb{E}[\delta \mid \hat{b}, \hat{r}] \\ &= (1 + \tau sp\rho_{b\delta|r})\eta_b + (sp\rho_{r\delta|b} - 1)\tau\eta_r + \varsigma_b + \tau sp\varsigma_\delta - \tau\varsigma_r. \end{aligned}$$

Suppose, as is likely to be the case, that $\rho_{b\delta|r} > 0$ (Lubowski *et al.* 2006). This means that the rent-conditioned partial correlation between environmental benefits per unit forgone output and the ILI is positive. Then the partially informed index should place more weight on environmental benefits than should the fully informed index in order to use b as a signal indicating a large δ value. The measure of environmental benefits needs to take on the weight of the correlation when no explicit measure of insurance subsidy costs is included in the index. Suppose too, as is also likely to be the case, that $\rho_{r\delta|b} < 0$ (Lubowski *et al.* 2006). Then the rent costs in the partially informed index should be weighed more heavily as a negative component than when insurance costs are included. This is because low-rent land suggests land with high insurance subsidy cost so that enticing low-rent land to enroll likely has a two-fold benefit.

Why This Design?

Two somewhat puzzling features of the CRP are (*a*) why land is enrolled entirely or not at all, rather than use an agro-environmental scheme whereby production occurs but some production

practices are proscribed in return for compensation (Fischer *et al.* 2008); and (*b*) why saved crop insurance subsidies are not accounted for. We ask whether these features have arisen from some recognition of problems that would arise were the situation otherwise. A review of the literature suggests that these questions may be related. Although the incentives design problem at hand, one of screening acres, does not involve moral hazard and is not really one of adverse selection, the principal-agent framework comes to mind.

Hölmstrom and Milgrom (1991) and Baker (2000, 2002) have used this approach to identify cases where it is better not to reward for readily measured dimensions of performance if another dimension is less readily measured. This is because one may get what can be measured well at the expense of losing what can be measured poorly. So perhaps it is that measures of subsidy losses are too noisy, or that inclusion of such measures would distort given the absence of some measures on important dimensions that are correlated.

That non-zero correlations exist between important dimensions to the services provided by retired land is clear enough. What is less clear is why this might motivate exclusion of a measure on insurance losses. It can hardly be that a measure on insurance losses is considered too noisy. The government has long histories of yield and loss performances at levels less aggregated than the county level, and often at the field level. Such data are used for rate-setting by crop insurance companies that provide the contracts that the government subsidizes. In addition, measures on environmental benefits (e.g., carbon sequestration) are included even though the scientific foundation for these measures is as yet far from complete.

This principal-agent literature also seeks to understand why outside activities on job time are sometimes proscribed rather than let output incentives implicitly penalize non-output activities. Hölmstrom and Milgrom (1991) argue that this can be because it is easier to prohibit an action entirely than to allow some of it but attempt to monitor and control the extent. There may be some merit to this reasoning when it is applied to, say, an upper bound on fertilizer or

pesticide application rates or use of no-till cultivation. But agro-environmental schemes that do precisely this have been in place in European Union countries for many years, apparently without widespread compliance problems (Baylis *et al.* 2008).

Whether these restrictions increase risk in the sense of the Lorenz order is not clear. Considerable research has come to ambiguous conclusions on whether more nitrogen generates a more variable distribution (Sheriff 2005) whereas it is almost certainly true that more nitrogen increases mean yield. In light of observations made concerning equation (5), if we assume that nitrogen has no effect on yield variability then empirical evidence tilts toward the hypothesis that more nitrogen induces a yield distribution that is less dispersed in the Lorenz order sense. And, as previously mentioned, the Just-Pope (1979) analysis of Day's data set suggests that nitrogen reduces risk in the Lorenz order sense for corn and oats.

Pesticides almost certainly increase mean yield. At first glance one may also posit that pesticides reduce yield variability in that they reduce the probability and extent of bad state outcomes while having no effect in good state outcomes. The situation would be similar to a lower contraction, as studied in Result 2. This would mean that pesticide use should reduce dispersion in the Lorenz order sense. Horowitz and Lichtenberg (1994), however, suggest that the matter is more complicated because overall crop growing conditions may correlate positively with the extent of the pest problem to be controlled. So the effect of pesticide use may more closely parallel Result 1 instead. Hurd (1994), for California cotton, and Shankar, Bennett and Morse (2008), for South African cotton, are among several works finding no significant effect of pesticides on yield variability. If pesticides increase mean yield and have no effect on variability then the ILI index will likely decline with an increase in their use.

On balance, we conclude that restrictions on the use of inputs that are protective in function would likely increase the cost of an insurance subsidy policy. It may be that, rather than restrict input use when producing on land that is marginal as cropland, is environmentally sensitive,

and has high yield variability given output level, it would be better to remove the land from production entirely. These arguments still do not explain why measures of insurance losses are excluded from the index.

Commencing with an assumed need in a political economy to compensate growers in the event of a poor harvest, Innes (2003) infers that any pre-harvest commitment not to do so is not credible. This is largely consistent with the history of U.S. federal disaster relief interventions. In his model, post-harvest intervention will take the form of revenue support. Yield takes the form of $\alpha_k \theta_k y_k$ where $\alpha_k > 0$, $y_k > 0$, and $\theta_k > 0$ are *k*th acre productivity, planned output or intensity decision, and multiplicative random component, respectively. There is no reference to environmental benefits from retiring land or from not farming it as intensively. Relative to efficient *laissez faire*, a positive probability of intervention when θ_k is low will draw marginal, but untilled, land into production. It will also reduce the incentive to farm land intensively.

To address these inefficiencies Innes suggests an integrated program of (*a*) revenue insurance with a 100% subsidy, since political intervention is inevitable; plus (*b*) coupled crop price support to offset the moral hazard effect arising from revenue insurance; plus (*c*) a participation fee to screen out (i.e., retire) inefficient low α_k acres that were drawn into production because of the state-contingent income support. The revenue insurance can also be viewed as crop insurance since price is non-stochastic in his model. The package can deliver first-best.

Innes notes that this policy package can also, at a stretch, motivate the CRP. The participation fee stick is not politically feasible. But it could be replaced by the carrot of a buyout program using an environmental criterion index since the environmental criteria are generally strongly correlated with low-productivity land. As he recognizes, this interpretation reaches outside his model, where the government's initial objective function places no weight on environmental objectives. But his insights are likely to be robust to any realistically

calibrated model allowing for environmental benefits.

Innes's well-crafted model sidesteps the issue of choosing acres to retire in light of social welfare losses due to insurance subsidy costs. This is because yield specification $\alpha_k \theta_k y_k$ has a constant ILI, regardless of acre chosen where this point was made under Result 5 above. Being a constant over acres in an index such as (3), it would be irrelevant. The issue we address is silent in Innes's model as the index used to sort land is unidimensional. But this is not reality, where a multitude of empirical studies attest that yield standard deviation per expected bushel produced varies over space, e.g., Kucharik and Ramankutty (2005).

Perhaps the most likely reason insurance subsidy losses are excluded in the EBI calculation is that the connection is not recognized in political calculations. Environmental and risk management programs have traditionally had different administrative structures and distinct parties recognized as interested. Perceptions on the need for policy innovations differ over time so that the items seldom enter the same debate.

Conclusion

The intent of this article is to promote a better understanding of the consequences of coherence among policies that affect land use. It shows how avoided crop insurance subsidy costs should be calculated for entry as a cost consideration in an environmental benefit index. The potential benefits of doing so are two-fold: to assist in managing the total cost of these programs and to better screen land into their best use. The Lorenz order was identified to best rank yield distributions over candidate acres for inclusion in a retirement program. Accurate index values for use at the acre level would be difficult to identify. But this is also true of environmental indices that are used. Empirical methods for comparing distributions in the Lorenz order sense are well developed (Bhattacharya 2007; Bresson 2009). These methods may provide practical guidance on how to modify land retirement enrollment procedures.

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Appendix

Demonstration of Result 1: The setting implies

$$\text{(A1)}\quad \begin{aligned} \delta_j(\beta) &= \frac{\pi \max[\beta\mu_j - y_j^l, 0] + (1-\pi)\max[\beta\mu_j - y_j^h, 0]}{\mu_j}; \\ \delta_k(\beta) &= \frac{\pi \max[\beta\mu_j + \beta(1-\pi)\phi - y_j^l, 0] + (1-\pi)\max[\beta\mu_j + \beta(1-\pi)\phi - \phi - y_j^h, 0]}{\mu_j + (1-\pi)\phi}. \end{aligned}$$

Note that $\beta \in [0,1]$ ensures $\beta\mu_j + \beta(1-\pi)\phi - \phi \le \beta\mu_j \le \beta\mu_j + \beta(1-\pi)\phi$. Write $A = [0, \beta\mu_j + \beta(1-\pi)\phi - \phi)$, $B = [\beta\mu_j + \beta(1-\pi)\phi - \phi, \beta\mu_j)$, $C = [\beta\mu_j, \beta\mu_j + \beta(1-\pi)\phi)$, and $D = [\beta\mu_j + \beta(1-\pi)\phi, \infty)$. There are many cases that need to be considered, some of which turn out to be impossible. The cases in which $y_j^h \in A \bigcup B$ are impossible since that would mean that $\mu_j < \beta\mu_j$, ruled out by $\beta \in [0,1]$. This leaves

Case 1: $y_j^l \in A \bigcup B$ and $y_j^h \in C \bigcup D$	Case 2: $y_j^l \in C$ and $y_j^h \in C \bigcup D$	Case 3: $y_j^l \in D$ and $y_j^h \in D$

<u>Case 1</u>: Here,

$$\text{(A2)}\quad \begin{aligned} \delta_j(\beta) &= \frac{\pi(\beta\mu_j - y_j^l)}{\mu_j} = \pi\beta - \frac{y_j^l}{\mu_j}; \\ \delta_k(\beta) &= \frac{\pi[\beta\mu_j + \beta(1-\pi)\phi - y_j^l]}{\mu_j + (1-\pi)\phi} = \pi\beta - \frac{\pi y_j^l}{\mu_j + (1-\pi)\phi} \ge \delta_j(\beta). \end{aligned}$$

<u>Case 2</u>: Here,

$$\text{(A3)}\quad \delta_j(\beta) = 0; \qquad \delta_k(\beta) = \frac{\pi[\beta\mu_j + \beta(1-\pi)\phi - y_j^l]}{\mu_j + (1-\pi)\phi} > 0.$$

<u>Case 3</u>: Here,

$$\text{(A4)}\quad \delta_j(\beta) = 0; \qquad \delta_k(\beta) = 0.$$

Thus, all cases have $\hat{\delta}_k(\beta) \geq \hat{\delta}_j(\beta)\ \forall \beta \in [0,1]$ as asserted.

Demonstration of Result 2: The setting implies

$$\text{(A5)}\quad \begin{aligned} \delta_j(\beta) &= \frac{\pi \max[\beta\mu_j - y_j^l, 0] + (1-\pi)\max[\beta\mu_j - y_j^h, 0]}{\mu_j}; \\ \delta_k(\beta) &= \frac{\pi \max[\beta\mu_j + \beta\pi\phi - \phi - y_j^l, 0] + (1-\pi)\max[\beta\mu_j + \beta\pi\phi - y_j^h, 0]}{\mu_j + \pi\phi}. \end{aligned}$$

It must be that $y_j^h = y_k^h \geq \beta\mu_j + \beta\pi\phi$ for otherwise we could not have $\mu_k = \mu_j + \pi\phi$, as required in light of the lower contraction attribute. And it must also be that $y_j^l < \mu_j + \pi\phi$ for otherwise $y_j^l + \phi > \mu_j + \pi\phi$ and $\mu_k > \mu_j + \pi\phi$, where $\mu_j + \pi\phi$ must equal μ_k by construction.

Write $E = [0, \beta\mu_j + \beta\pi\phi - \phi)$, $F = [\beta\mu_j + \beta\pi\phi - \phi, \beta\mu_j)$, $G = [\beta\mu_j, \beta\mu_j + \beta\pi\phi)$, and $H = [\beta\mu_j + \beta\pi\phi, \infty)$ where $y_j^h \in H$ and $y_j^l \notin H$. The cases we need to consider are

Case I: $y_j^l \in E$ and $y_j^h \in H$	Case II: $y_j^l \in F$ and $y_j^h \in H$	Case III: $y_j^l \in G$ and $y_j^h \in H$

<u>Case I</u>: Here,

$$\text{(A6)}\quad \begin{aligned} \delta_j(\beta) &= \frac{\pi(\beta\mu_j - y_j^l)}{\mu_j} = \pi\beta - \frac{\pi y_j^l}{\mu_j}; \\ \delta_k(\beta) &= \frac{\pi(\beta\mu_j + \beta\pi\phi - \phi - y_j^l)}{\mu_j + \pi\phi} = \pi\beta - \frac{\pi(\phi + y_j^l)}{\mu_j + \pi\phi} \\ &= \delta_j(\beta) + \frac{\pi y_j^l}{\mu_j} - \frac{\pi(\phi + y_j^l)}{\mu_j + \pi\phi} = \delta_j(\beta) + \frac{\pi\phi(y_j^l\pi - \mu_j)}{\mu_j(\mu_j + \pi\phi)} = \delta_j(\beta) - \frac{\pi\phi y_j^h(1-\pi)}{\mu_j(\mu_j + \pi\phi)} < \delta_j(\beta). \end{aligned}$$

<u>Case II</u>: Here,

$$\text{(A7)}\quad \delta_j(\beta) = \frac{\pi(\beta\mu_j - y_j^l)}{\mu_j} = \pi\beta - \frac{\pi y_j^l}{\mu_j} > 0; \qquad \delta_k(\beta) = 0.$$

<u>Case III</u>: Here,

(A8) $\delta_j(\beta) = 0; \qquad \delta_k(\beta) = 0.$

Thus $\hat{\delta}_k(\beta) \le \hat{\delta}_j(\beta)\ \forall \beta \in [0,1]$ in all cases, as asserted.

Demonstration of Result 3: This follows immediately from (5). Since $\mu_k = \mu_j$, the denominator does not change. Since $\beta\mu_k = \beta\mu_j$ in the numerator, the mean-preserving spread increases the expected value of the convex $\max[\cdot, 0]$ function in the numerator.

Demonstration of Result 4: We have $\mu_j = \pi\mu_k$ so that independence secures

$$\text{(A9)} \quad \begin{aligned} \hat{\delta}_k(\beta) &= \frac{\mathbb{E}[\max(\beta\pi\mu_j - y_k, 0)]}{\pi\mu_j} = \frac{\pi\mathbb{E}[\max(\beta\pi\mu_j - y_j, 0)] + (1-\pi)\mathbb{E}[\max(\beta\pi\mu_j, 0)]}{\pi\mu_j} \\ &= \frac{\mathbb{E}[\max(\beta\mu_j - y_j, (1-\pi)\beta\mu_j)]}{\mu_j} \ge \frac{\mathbb{E}[\max(\beta\mu_j - y_j, 0)]}{\mu_j} = \hat{\delta}_j(\beta). \end{aligned}$$