Let's Take the Con out of Econometrics

By EDWARD E. LEAMER*

Econometricians would like to project the image of agricultural experimenters who divide a farm into a set of smaller plots of land and who select randomly the level of fertilizer to be used on each plot. If some plots are assigned a certain amount of fertilizer while others are assigned none, then the difference between the mean yield of the fertilized plots and the mean yield of the unfertilized plots is a measure of the effect of fertilizer on agricultural yields. The econometrician's humble job is only to determine if that difference is large enough to suggest a real effect of fertilizer, or is so small that it is more likely due to random variation.

This image of the applied econometrician's art is grossly misleading. I would like to suggest a more accurate one. The applied econometrician is like a farmer who notices that the yield is somewhat higher under trees where birds roost, and he uses this as evidence that bird droppings increase yields. However, when he presents this finding at the annual meeting of the American Ecological Association, another farmer in the audience objects that he used the same data but came up with the conclusion that moderate amounts of shade increase yields. A bright chap in the back of the room then observes that these two hypotheses are indistinguishable, given the available data. He mentions the phrase "identification problem," which, though no one knows quite what he means, is said with such authority that it is totally convincing. The meeting reconvenes in the halls and in the bars, with heated discussion whether this is the kind of work that merits promotion from Associate to Full Farmer; the Luminists strongly opposed to promotion and the Aviophiles equally strong in favor.

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One should not jump to the conclusion that there is necessarily a substantive difference between drawing inferences from experimental as opposed to nonexperimental data. The images I have drawn are deliberately prejudicial. First, we had the experimental scientist with hair neatly combed, wide eyes peering out of horn-rimmed glasses, a white coat, and an electronic calculator for generating the random assignment of fertilizer treatment to plots of land. This seems to contrast sharply with the nonexperimental farmer with overalls, unkempt hair, and bird droppings on his boots. Another image, drawn by Orcutt, is even more damaging: "Doing econometrics is like trying to learn the laws of electricity by playing the radio." However, we need not now submit to the tyranny of images, as many of us have in the past.

I. Is Randomization Essential?

What is the real difference between these two settings? Randomization seems to be the answer. In the experimental setting, the fertilizer treatment is "randomly" assigned to plots of land, whereas in the other case nature did the assignment. Now it is the tyranny of words that we must resist. "Random" does not mean adequately mixed in *every* sample. It only means that on the average, the fertilizer treatments are adequately mixed. Randomization implies that the least squares estimator is "unbiased," but that definitely does not mean that for each sample the estimate is correct. Sometimes the estimate is too high, sometimes too low. I am reminded of the lawyer who remarked that "when I was a young man I lost many cases that I should have won, but when I grew older I won many that I should have lost, so on the average justice was done."

In particular, it is possible for the randomized assignment to lead to exactly the same allocation as the nonrandom assignment,

namely, with treated plots of land all being under trees and with nontreated plots of land all being away from trees. I submit that, if this is the outcome of the randomization, then the randomized experiment and the nonrandomized experiment are exactly the same. Many econometricians would insist that there is a difference, because the randomized experiment generates "unbiased" estimates. But all this means is that, if this particular experiment yields a gross overestimate, some other experiment yields a gross underestimate.

Randomization thus does not assure that each and every experiment is "adequately mixed," but randomization does make "adequate mixing" probable. In order to make clear what I believe to be the true value of randomization, let me refer to the model

$$(1) \qquad Y_i = \alpha + \beta F_i + \gamma L_i + U_i,$$

where Y_i is the yield of plot i; F_i is the fertilizer assigned to plot i; L_i is the light falling on plot i; U_i is the unspecified influence on the yield of plot i, and where β, the fertilizer effect, is the object of the inferential exercise. We may suppose to begin the argument that the light level is expensive to measure and that it is decided to base an estimate of β initially only on measurement of Y_i and F_i. We may assume also that the natural experiment produces values for F_i, L_i, and U_i with expected values $E(U_i|F_i) = 0$ and $E(L_i|F_i) = r_0 + r_1 F_i$. In the more familiar parlance, it is assumed that the fertilizer level and the residual effects are uncorrelated, but the fertilizer level and the light level are possibly correlated. As every beginning econometrics student knows, if you omit from a model a variable which is correlated with included variables, bad things happen. These bad things are revealed to the econometrician by computing the conditional mean of Y given F but not L:

$$(2) \quad \begin{aligned} E(Y|F) &= \alpha + \beta F + \gamma E(L|F) \\ &= \alpha + \beta F + \gamma(r_0 + r_1 F) \\ &\equiv (\alpha + \alpha^*) + (\beta + \beta^*) F, \end{aligned}$$

where $\alpha^* = \gamma r_0$ and $\beta^* = \gamma r_1$. The linear regression of Y on F provides estimates of the parameters of the conditional distribution of Y given F, and in this case the regression coefficients are estimates not of α and β, but rather of $\alpha + \alpha^*$ and $\beta + \beta^*$. The parameters α^* and β^* measure the bias in the least squares estimates. This bias could be due to left-out variables, or to measurement errors in F, or to simultaneity.

When observing a nonexperiment, the bias parameters α^* and β^* can be thought to be small, but they cannot sensibly be treated as exact zeroes. The notion that the bias parameters are small can be captured by the assumption that α^* and β^* are drawn from a normal distribution with zero means and covariance matrix M. The model can then be written as $Y = \alpha + \beta F + \varepsilon$, where ε is the sum of three random variables: $U + \alpha^* + \beta^* F$. Because the error term ε is not spherical, the proper way to estimate α and β is generalized least squares. My 1974 article demonstrates that if (a, b) represent the least squares estimates of (α, β), then the generalized least squares estimates $(\hat{\alpha}, \hat{\beta})$ are also equal to (a, b):

$$(3) \qquad \begin{pmatrix} \hat{\alpha} \\ \hat{\beta} \end{pmatrix} = \begin{pmatrix} a \\ b \end{pmatrix},$$

and if S represents the sample covariance matrix for the least squares estimates, then the sample covariance matrix for $(\hat{\alpha}, \hat{\beta})$ is

$$(4) \qquad Var(\hat{\alpha}, \hat{\beta}) = S + M,$$

where M is the covariance matrix of (α^*, β^*).

The meaning of equation (3) is that unless one knows the direction of the bias, the possibility of bias does not call for any adjustment to the estimates. The possibility of bias does require an adjustment to the covariance matrix (4). The uncertainty is composed of two parts: the usual sampling uncertainty S plus the misspecification uncertainty M. As sample size grows, the sampling uncertainty S ever decreases, but the misspecification uncertainty M remains ever constant. The misspecification matrix M that we must add to the least squares variance

matrix is just the (prior) variance of the bias coefficients (α^*, β^*). If this variance matrix is small, the least squares bias is likely to be small. If M is large, it is correspondingly probable that (α^*, β^*) is large.

It would be a remarkable bootstrap if we could determine the extent of the misspecification from the data. The data in fact contain no information about the size of the bias, a point which is revealed by studying the likelihood function. The misspecification matrix M is therefore a pure prior concept. One must decide independent of the data how good the nonexperiment is.

The formal difference between a randomized experiment and a natural experiment is measured by the matrix M. If the treatment is randomized, the bias parameters (α^*, β^*) are exactly zero, or, equivalently, the matrix M is a zero matrix. If M is zero, the least squares estimates are consistent. If M is not zero, as in the natural experiment, there remains a fixed amount of specification uncertainty, independent of sample size.

There is therefore a sharp difference between inference from randomized experiments and inference from natural experiments. This seems to draw a sharp distinction between economics where randomized experiments are rare and "science" where experiments are routinely done. But the fact of the matter is that no one has ever designed an experiment that is free of bias, and no one can. As it turns out, the technician who was assigning fertilizer levels to plots of land, took his calculator into the fields, and when he was out in the sun, the calculator got heated up and generated large "random" numbers, which the technician took to mean no fertilizer; and when he stood under the shade of the trees, his cool calculator produced small numbers, and these plots received fertilizer.

You may object that this story is rather fanciful, but I need only make you think it is possible, to force you to set $M \neq 0$. Or if you think a computer can really produce random numbers (calculated by a mathematical formula and therefore perfectly predictable!), I will bring up mismeasurement of the fertilizer level, or human error in carrying out the computer instructions. Thus, the attempt to randomize and the attempt to measure accurately ensures that M is small, but not zero, and the difference between scientific experiments and natural experiments is difference in degree, but not in kind. Admittedly however, the misspecification uncertainty in many experimental settings may be so small that it is well approximated by zero. This can very rarely be said in nonexperimental settings.

Examples may be ultimately convincing. There is a great deal of empirical knowledge in the science of astronomy, yet there are no experiments. Medical knowledge is another good example. I was struck by a headline in the January 5, 1982 *New York Times*: "Life Saving Benefits of Low-Cholesterol Diet Affirmed in *Rigorous* Study." The article describes a randomized experiment with a control group and a treated group. "Rigorous" is therefore interpreted as "randomized." As a matter of fact, there was a great deal of evidence suggesting a link between heart disease and diet before any experiments were performed on humans. There were cross-cultural comparisons and there were animal studies. Actually, the only reason for performing the randomized experiment was that someone believed there was pretty clear nonexperimental evidence to begin with. The nonexperimental evidence was, of course, inconclusive, which in my language means that the misspecification uncertainty M remained uncomfortably large. The fact that the Japanese have both less incidence of heart disease and also diets lower in cholesterol compared to Americans is not convincing evidence, because there are so many other factors that remain unaccounted for. The fact that pigs on a high cholesterol diet develop occluded arteries is also not convincing, because the similarity in physiology in pigs and humans can be questioned.

When the sampling uncertainty S gets small compared to the misspecification uncertainty M, it is time to look for other forms of evidence, experiments or nonexperiments. Suppose I am interested in measuring the width of a coin, and I provide rulers to a room of volunteers. After each volunteer has reported a measurement, I compute the mean and standard deviation, and I conclude that

the coin has width 1.325 millimeters with a standard error of .013. Since this amount of uncertainty is not to my liking, I propose to find three other rooms full of volunteers, thereby multiplying the sample size by four, and dividing the standard error in half. That is a silly way to get a more accurate measurement, because I have already reached the point where the sampling uncertainty S is very small compared with the misspecification uncertainty M. If I want to increase the true accuracy of my estimate, it is time for me to consider using a micrometer. So too in the case of diet and heart disease. Medical researchers had more or less exhausted the vein of nonexperimental evidence, and it became time to switch to the more expensive but richer vein of experimental evidence.

In economics, too, we are switching to experimental evidence. There are the laboratory experiments of Charles Plott and Vernon Smith (1978) and Smith (1980), and there are the field experiments such as the Seattle/Denver income maintenance experiment. Another way to limit the misspecification error M is to gather different kinds of nonexperiments. Formally speaking, we will say that experiment 1 is qualitatively different from experiment 2 if the bias parameters (α_1^*, β_1^*) are distributed independently of the bias parameters (α_2^*, β_2^*). In that event, simple averaging of the data from the two experiments yields average bias parameters $(\alpha_1^* + \alpha_2^*,\ \beta_1^* + \beta_2^*)/2$ with misspecification variance matrix $M/2$, half as large as the (common) individual variances. Milton Friedman's study of the permanent income hypothesis is the best example of this that I know. Other examples are hard to come by. I believe we need to put much more effort into identifying qualitatively different and convincing kinds of evidence.

Parenthetically, I note that traditional econometric theory, which does not admit experimental bias, as a consequence also admits no "hard core" propositions. Demand curves can be shown to be positively sloped. Utility can be shown not to be maximized. Econometric evidence of a positively sloped demand curve would, as a matter of fact, be routinely explained in terms of simultaneity bias. If utility seems not to have been maximized, it is only that the econometrician has misspecified the utility function. The misspecification matrix M thus forms Imre Lakatos' "protective belt" which protects certain hard core propositions from falsification.

II. Is Control Essential?

The experimental scientist who notices that the fertilizer treatment is correlated with the light level can correct his experimental design. He can control the light level, or he can allocate the fertilizer treatment in such a way that the fertilizer level and the light level are not perfectly correlated.

The nonexperimental scientist by definition cannot control the levels of extraneous influences such as light. But he can control for the variable light level by including light in the estimating equation. Provided nature does not select values for light and values for fertilizer levels that are perfectly correlated, the effect of fertilizer on yields can be estimated with a multiple regression. The collinearity in naturally selected treatment variables may mean that the data evidence is weak, but it does not invalidate in any way the usual least squares estimates. Here, again, there is no essential difference between experimental and nonexperimental inference.

III. Are the Degrees of Freedom Inadequate with Nonexperimental Data?

As a substitute for experimental control, the nonexperimental researcher is obligated to include in the regression equation all variables that might have an important effect. The NBER data banks contain time-series data on 2,000 macroeconomic variables. A model explaining gross national product in terms of all these variables would face a severe degrees-of-freedom deficit since the number of annual observations is less than thirty. Though the number of observations of any phenomenon is clearly limited, the number of explanatory variables is logically unlimited. If a polynomial could have a degree as high as k, it would usually be admitted that the degree could be $k+1$ as well. A theory that allows k lagged explanatory vari-

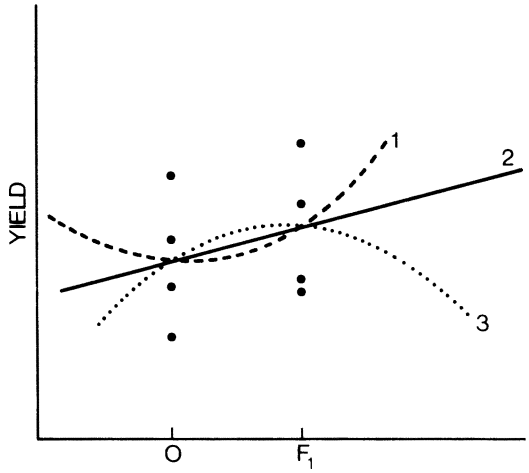

FIGURE 1. HYPOTHETICAL DATA AND THREE ESTIMATED QUADRATIC FUNCTIONS

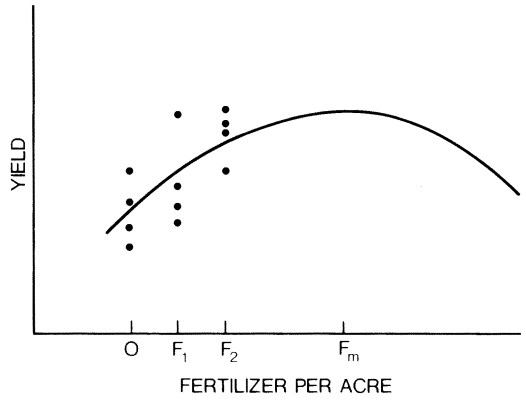

FIGURE 2. HYPOTHETICAL DATA AND ESTIMATED QUADRATIC FUNCTION

ables would ordinarily allow $k+1$. If the level of money might affect *GNP*, then why not the number of presidential sneezes, or the size of the polar ice cap?

The number of explanatory variables is unlimited in a nonexperimental setting, but it is also unlimited in an experimental setting. Consider again the fertilizer example in which the farmer randomly decides either to apply F_1 pounds of fertilizer per acre or zero pounds, and obtains the data illustrated in Figure 1. These data admit the inference that fertilizer level F_1 produces higher yields than no fertilizer. But the farmer is interested in selecting the fertilizer level that maximizes profits. If it is hypothesized that yield is a linear function of the fertilizer intensity $Y = \alpha + \beta F + U$, then profits are

$$Profits = pA(\alpha + \beta F + U) - p_F AF,$$

where A is total acreage, p is the product price, and p_F is the price per pound of fertilizer. This profit function is linear in F with slope $A(\beta p - p_F)$. The farmer maximizes profits therefore by using no fertilizer if the price of fertilizer is high, $\beta p < p_F$, and using an unlimited amount of fertilizer if the price is low, $\beta p > p_F$. It is to be expected that you will find this answer unacceptable for one of several reasons:

1) When the farmer tries to buy an unlimited amount of fertilizer, he will drive up its price, and the problem should be reformulated to make p_F a function of F.

2) Uncertainty in the fertilizer effect β causes uncertainty in profits, $Variance(profits) = p^2A^2F^2Var(\beta)$, and risk aversion will limit the level of fertilizer applied.

3) The yield function is nonlinear.

Economic theorists doubtless find reasons 1) and 2) compelling, but I suspect that the real reason farmers don't use huge amounts of fertilizer is that the marginal increase in the yield eventually decreases. Plants don't grow in fertilizer alone.

So let us suppose that yield is a quadratic function of fertilizer intensity, $Y = \alpha + \beta_1 F + \beta_2 F^2 + U$, and suppose we have only the data illustrated in Figure 1. Unfortunately, there are an infinite number of quadratic functions all of which fit the data equally well, three of which are drawn. If there were no other information available, we could conclude only that the yield is higher at F_1 than at zero. Formally speaking, there is an identification problem, which can be solved by altering the experimental design. The yield must be observed at a third point, as in Figure 2, where I have drawn the least squares estimated quadratic function and have indicated the fertilizer intensity F_m that maximizes the yield. I expect that most people would question whether these data admit the

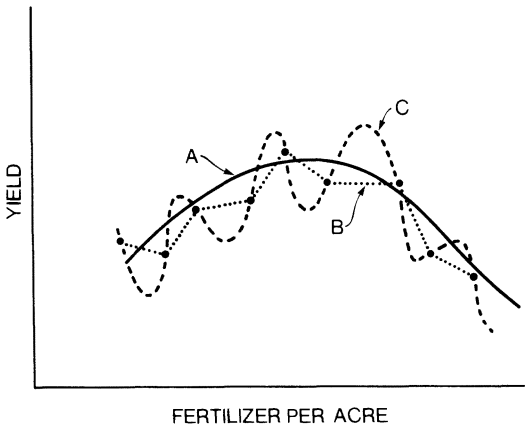

FIGURE 3. HYPOTHETICAL DATA AND THREE ESTIMATED FUNCTIONS

inference that the yield is maximized at F_m. Actually, after inspection of this figure, I don't think anything can be inferred except that the yield at F_2 is higher than at F_1, which in turn is higher than at zero. Thus I don't believe the function is quadratic. If it is allowed to be a cubic then again there is an identification problem.

This kind of logic can be extended indefinitely. One can always find a set of observations that will make the inferences implied by a polynomial of degree p seem silly. This is true regardless of the degree p. Thus no model with a finite number of parameters is actually believed, whether the data are experimental or nonexperimental.

IV. Do We Need Prior Information?

A model with an infinite number of parameters will allow inference from a finite data set only if there is some prior information that effectively constrains the ranges of the parameters. Figure 3 depicts another hypothetical sequence of observations and three estimated relationships between yield and fertilizer. I believe the solid line A is a better representation of the relationship than either of the other two. The piecewise linear form B fits the data better, but I think this peculiar meandering function is highly unlikely on an a priori basis. Though B and C fit the data equally well, I believe that B is much more likely than C. What I am revealing is the a priori opinion that the function is likely to be smooth and single peaked.

What should now be clear is that data alone cannot reveal the relationship between yield and fertilizer intensity. Data can reveal the yield at sampled values of fertilizer intensities, but in order to interpolate between these sampled values, we must resort to subjective prior information.

Economists have inherited from the physical sciences the myth that scientific inference is objective, and free of personal prejudice. This is utter nonsense. All knowledge is human belief; more accurately, human opinion. What often happens in the physical sciences is that there is a high degree of conformity of opinion. When this occurs, the opinion held by most is asserted to be an objective fact, and those who doubt it are labelled "nuts." But history is replete with examples of opinions losing majority status, with once-objective "truths" shrinking into the dark corners of social intercourse. To give a trivial example, coming now from California I am unsure whether fat ties or thin ties are aesthetically more pleasing.

The false idol of objectivity has done great damage to economic science. Theoretical econometricians have interpreted scientific objectivity to mean that an economist must identify exactly the variables in the model, the functional form, and the distribution of the errors. Given these assumptions, and given a data set, the econometric method produces an objective inference from a data set, unencumbered by the subjective opinions of the researcher.

This advice could be treated as ludicrous, except that it fills all the econometric textbooks. Fortunately, it is ignored by applied econometricians. The econometric art as it is practiced at the computer terminal involves fitting many, perhaps thousands, of statistical models. One or several that the researcher finds pleasing are selected for reporting purposes. This searching for a model is often well intentioned, but there can be no doubt that such a specification search invalidates the traditional theories of inference. The concepts of unbiasedness, consistency, efficiency, maximum-likelihood estimation,

in fact, all the concepts of traditional theory, utterly lose their meaning by the time an applied researcher pulls from the bramble of computer output the one thorn of a model he likes best, the one he chooses to portray as a rose. The consuming public is hardly fooled by this chicanery. The econometrician's shabby art is humorously and disparagingly labelled "data mining," "fishing," "grubbing," "number crunching." A joke evokes the Inquisition: "If you torture the data long enough, Nature will confess" (Coase). Another suggests methodological fickleness: "Econometricians, like artists, tend to fall in love with their models" (wag unknown). Or how about: "There are two things you are better off not watching in the making: sausages and econometric estimates."

This is a sad and decidedly unscientific state of affairs we find ourselves in. Hardly anyone takes data analyses seriously. Or perhaps more accurately, hardly anyone takes anyone else's data analyses seriously. Like elaborately plumed birds who have long since lost the ability to procreate but not the desire, we preen and strut and display our t-values.

If we want to make progress, the first step we must take is to discard the counterproductive goal of objective inference. The dictionary defines an inference as a logical conclusion based on a set of facts. The "facts" used for statistical inference about θ are first the data, symbolized by x, second a conditional probability density, known as a sampling distribution, $f(x|\theta)$, and, third, explicitly for a Bayesian and implicitly for "all others," a marginal or prior probability density function $f(\theta)$. Because both the sampling distribution and the prior distribution are actually *opinions* and not *facts*, a statistical inference is and must forever remain an *opinion*.

What is a fact? A fact is merely an opinion held by all, or at least held by a set of people you regard to be a close approximation to all.[1] For some that set includes only one person. I myself have the opinion that Andrew Jackson was the sixteenth president of the United States. If many of my friends agree, I may take it to be a fact. Actually, I am most likely to regard it to be a fact if the authors of one or more books say it is so.

The difference between a fact and an opinion for purposes of decision making and inference is that when I use opinions, I get uncomfortable. I am not too uncomfortable with the opinion that error terms are normally distributed because most econometricians make use of that assumption. This observation has deluded me into thinking that the opinion that error terms are normal may be a fact, when I know deep inside that normal distributions are actually used only for convenience. In contrast, I am *quite* uncomfortable using a prior distribution, mostly I suspect because hardly anyone uses them. If convenient prior distributions were used as often as convenient sampling distributions, I suspect that I could be as easily deluded into thinking that prior distributions are facts as I have been into thinking that sampling distributions are facts.

To emphasize this hierarchy of statements, I display them in order: truths; facts; opinions; conventions. Note that I have added to the top of the order, the category truths. This will appeal to those of you who feel compelled to believe in such things. At the bottom are conventions. In practice, it may be difficult to distinguish a fact from a convention, but when facts are clearly unavailable, we must strongly resist the deceit or delusion that conventions can represent.

What troubles me about using opinions is their whimsical nature. Some mornings when I arise, I have the opinion that Raisin Bran is better than eggs. By the time I get to the kitchen, I may well decide on eggs, or oatmeal. I usually do recall that the sixteenth president distinguished himself. Sometimes I think he was Jackson; often I think he was Lincoln.

A data analysis is similar. Sometimes I take the error terms to be correlated, sometimes uncorrelated; sometimes normal and sometimes nonnormal; sometimes I include observations from the decade of the fifties, sometimes I exclude them; sometimes the

[1] This notion of "truth by consensus" is espoused by Thomas Kuhn (1962) and Michael Polanyi (1964). Oscar Wilde agrees by dissent: "A truth ceases to be true when more than one person believes it."

equation is linear and sometimes nonlinear; sometimes I control for variable z, sometimes I don't. Does it depend on what I had for breakfast?

As I see it, the fundamental problem facing econometrics is how adequately to control the whimsical character of inference, how sensibly to base inferences on opinions when facts are unavailable. At least a partial solution to this problem has already been formed by practicing econometricians. A common reporting style is to record the inferences implied by alternative sets of opinions. It is not unusual to find tables that show how an inference changes as variables are added to or deleted from the equation. This kind of sensitivity analysis reports special features of the mapping from the space of assumptions to the space of inferences. The defect of this style is that the coverage of assumptions is infinitesimal, in fact a zero volume set in the space of assumptions. What is needed instead is a more complete, but still economical way to report the mapping of assumptions into inferences. What I propose to do is to develop a correspondence between regions in the assumption space and regions in the inference space. I will report that all assumptions in a certain set lead to essentially the same inference. Or I will report that there are assumptions within the set under consideration that lead to radically different inferences. In the latter case, I will suspend inference and decision, or I will work harder to narrow the set of assumptions.

Thus what I am asserting is that the choice of a particular sampling distribution, or a particular prior distribution, is inherently whimsical. But statements such as "The sampling distribution is symmetric and unimodal" and "My prior is located at the origin" are not necessarily whimsical, and in certain circumstances do not make me uncomfortable.

To put this somewhat differently, an inference is not believable if it is fragile, if it can be reversed by minor changes in assumptions. As consumers of research, we correctly reserve judgment on an inference until it stands up to a study of fragility, usually by other researchers advocating opposite opinions. It is, however, much more efficient for individual researchers to perform their own sensitivity analyses, and we ought to be demanding much more complete and more honest reporting of the fragility of claimed inferences.

The job of a researcher is then to report economically and informatively the mapping from assumptions into inferences. In a slogan, "The mapping is the message." The mapping does not depend on opinions (assumptions), but reporting the mapping economically and informatively does. A researcher has to decide which assumptions or which sets of alternative assumptions are worth reporting. A researcher is therefore forced either to anticipate the opinions of his consuming public, or to recommend his own opinions. It is actually a good idea to do both, and a serious defect of current practice is that it concentrates excessively on convincing one's self and, as a consequence, fails to convince the general professional audience.

The whimsical character of econometric inference has been partially controlled in the past by an incomplete sensitivity analysis. It has also been controlled by the use of conventions. The normal distribution is now so common that there is nothing at all whimsical in its use. In some areas of study, the list of variables is partially conventional, often based on whatever list the first researcher happened to select. Even conventional prior distributions have been proposed and are used with nonnegligible frequency. I am referring to Robert Shiller's (1973) smoothness prior for distributed lag analysis and to Arthur Hoerl and Robert Kennard's (1970) ridge regression prior. It used to aggravate me that these methods seem to find public favor whereas overt and complete Bayesian methods such as my own proposals (1972) for distributed lag priors are generally ignored. However, there is a very good reason for this: the attempt to form a prior distribution from scratch involves an untold number of partly arbitrary decisions. The public is rightfully resistant to the whimsical inferences which result, but at the same time is receptive to the use of priors in ways that control the whimsy. Though the use of conventions does control the whimsy, it can do so at the cost of relevance. Inferences based

on Hoerl and Kennard's conventional "ridge regression" prior are usually irrelevant, because it is rarely sensible to take the prior to be spherical and located at the origin, and because a closer approximation to prior belief can be suspected to lead to substantially different inferences. In contrast, the conventional assumption of normality at least uses a distribution which usually cannot be ruled out altogether. Still, we may properly demand a demonstration that the inferences are insensitive to this distributional assumption.

A. *The Horizon Problem: Sherlock Holmes Inference*

Conventions are not to be ruled out altogether, however. One can go mad trying to report completely the mapping from assumptions into inferences since the space of assumptions is infinite dimensional. A formal statistical analysis therefore has to be done within the limits of a reasonable horizon. An informed convention can usefully limit this horizon. If it turned out that sensible neighborhoods of distributions around the normal distribution 99 times out of 100 produced the same inference, then we could all agree that there are other more important things to worry about, and we may properly adopt the convention of normality. The consistency of least squares estimates under wide sets of assumptions is used improperly as support for this convention, since the inferences from a given finite sample may nonetheless be quite sensitive to the normality assumption.[2]

The truly sharp distinction between inference from experimental and inference from nonexperimental data is that experimental inference sensibly admits a conventional horizon in a critical dimension, namely the choice of explanatory variables. If fertilizer is randomly assigned to plots of land, it is conventional to restrict attention to the relationship between yield and fertilizer, and to proceed as if the model were perfectly specified, which in my notation means that the misspecification matrix M is the zero matrix. There is only a small risk that when you present your findings, someone will object that fertilizer and light level are correlated, and there is an even smaller risk that the conventional zero value for M will lead to inappropriate inferences. In contrast, it would be foolhardy to adopt such a limited horizon with nonexperimental data. But if you decide to include light level in your horizon, then why not rainfall; and if rainfall, then why not temperature; and if temperature, then why not soil depth, and if soil depth, then why not the soil grade; ad infinitum. Though this list is never ending, it can be made so long that a nonexperimental researcher can feel as comfortable as an experimental researcher that the risk of having his findings upset by an extension of the horizon is very low. The exact point where the list is terminated must be whimsical, but the inferences can be expected not to be sensitive to the termination point if the horizon is wide enough.

Still, the horizon within which we all do our statistical analyses has to be ultimately troublesome, since there is no formal way to know what inferential monsters lurk beyond our immediate field of vision. "Diagnostic" tests with explicit alternative hypotheses such as the Durbin-Watson test for first-order autocorrelation do not truly ask if the horizon should be extended, since first-order autocorrelation is explicitly identified and clearly in our field of vision. Diagnostic tests such as goodness-of-fit tests, without explicit alternative hypotheses, are useless since, if the sample size is large enough, any maintained hypothesis will be rejected (for example, no observed distribution is exactly normal). Such tests therefore degenerate into elaborate rituals for measuring the effective sample size.

The only way I know to ask the question whether the horizon is wide enough is to study the anomalies of the data. In the words of the physiologist, C. Bernard:

> A great surgeon performs operations for stones by a single method; later he

[2] In particular, least squares estimates are completely sensitive to the independence assumption, since by choice of sample covariance matrix a generalized least squares estimate can be made to assume any value whatsoever (see my 1981 paper).

> makes a statistical summary of deaths and recoveries, and he concludes from these statistics that the mortality law for this operation is two out of five. Well, I say that this ratio means literally nothing scientifically, and gives no certainty in performing the next operation. What really should be done, instead of gathering facts empirically, is to study them more accurately, each in its special determinism... by statistics, we get a conjecture of greater or less probability about a given case, but never any certainty, never any absolute determinism... only basing itself on experimental determinism can medicine become a true science.
>
> [1927, pp. 137–38]

A study of the anomalies of the data is what I have called "Sherlock Holmes" inference, since Holmes turns statistical inference on its head: "It is a capital mistake to theorize before you have all the evidence. It biases the judgements." Statistical theory counsels us to begin with an elicitation of opinions about the sampling process and its parameters; the theory, in other words. After that, data may be studied in a purely mechanical way. Holmes warns that this biases the judgements, meaning that a theory constructed before seeing the facts can be disastrously inappropriate and psychologically difficult to discard. But if theories are constructed after having studied the data, it is difficult to establish by how much, if at all, the data favor the data-instigated hypothesis. For example, suppose I think that a certain coefficient ought to be positive, and my reaction to the anomalous result of a negative estimate is to find another variable to include in the equation so that the estimate is positive. Have I found evidence that the coefficient is positive? It would seem that we should require evidence that is more convincing than the traditional standard. I have proposed a method for discounting such evidence (1974). Initially, when you regress yield on fertilizer as in equation (2), you are required to assess a prior distribution for the experimental bias parameter β^*; that is, you must select the misspecification matrix M. Then, when the least squares estimate of β turns out to be negative, and you decide to include in the equation the light level as well as the fertilizer level, you are obligated to form a prior for the light coefficient γ consistent with the prior for β^*, given that $\beta^* = \gamma r_1$, where r_1 is the regression coefficient of light on fertilizer.[3]

This method for discounting the output of exploratory data analysis requires a discipline that is lacking even in its author. It is consequently important that we reduce the risk of Holmesian discoveries by extending the horizon reasonably far. The degree of a polynomial or the order of a distributed lag need not be data instigated, since the horizon is easily extended to include high degrees and high orders. It is similarly wise to ask yourself before examining the data what you would do if the estimate of your favorite coefficient had the wrong sign. If that makes you think of a specific left-out variable, it is better to include it from the beginning.

Though it is wise to select a wide horizon to reduce the risk of Holmesian discoveries, it is mistaken then to analyze a data set as if the horizon were wide enough. Within the limits of a horizon, no revolutionary inference can be made, since all possible inferences are predicted in advance (admittedly, some with low probabilities). Within the horizon, inference and decision can be turned over completely to a computer. But the great human revolutionary discoveries are made when the horizon is extended for reasons that cannot be predicted in advance and cannot be computerized. If you wish to make such discoveries, you will have to poke at the horizon, and poke again.

V. An Example

This rhetoric is understandably tiring. Methodology, like sex, is better demonstrated than discussed, though often better anticipated than experienced. Accordingly, let me give you an example of what all this

[3] In a randomized experiment with $r_1 = 0$, the constraint $\beta^* = \gamma r_1$ is irrelevant, and you are free to play these exploratory games without penalty. This is a very critical difference between randomized experiments and nonrandomized nonexperiments.

ranting and raving is about. I trust you will find it even better in the experience than in the anticipation. A problem of considerable policy importance is whether or not to have capital punishment. If capital punishment had no deterrent value, most of us would prefer not to impose such an irreversible punishment, though, for a significant minority, the pure joy of vengeance is reason enough. The deterrent value of capital punishment is, of course, an empirical issue. The unresolved debate over its effectiveness began when evolution was judging the survival value of the vengeance gene. Nature was unable to make a decisive judgment. Possibly econometricians can.

In Table 1, you will find a list of variables that are hypothesized to influence the murder rate.[4] The data to be examined are state-by-state murder rates in 1950. The variables are divided into three sets. There are four deterrent variables that characterize the criminal justice system, or in economic parlance, the expected out-of-pocket cost of crime. There are four economic variables that measure the opportunity cost of crime. And there are four social/environmental variables that possibly condition the taste for crime. This leaves unmeasured only the expected rewards for criminal behavior, though these are possibly related to the economic and social variables and are otherwise assumed not to vary from state to state.

A simple regression of the murder rate on all these variables leads to the conclusion that each additional execution deters thirteen murders, with a standard error of seven. That seems like such a healthy rate of return, we might want just to randomly draft executees from the population at large. This proposal would be unlikely to withstand the scrutiny of any macroeconomists who are skilled at finding rational expectations equlibria.

The issue I would like to address instead is whether this conclusion is fragile or not. Does it hold up if the list of variables in the model is changed? Individuals with different experiences and different training will find different subsets of the variables to be candidates for omission from the equation. Five different lists of doubtful variables are reported in Table 2. A right winger expects

TABLE 1—VARIABLES USED IN THE ANALYSIS

a. Dependent Variable
   M = Murder rate per 100,000, FBI estimate.
b. Independent Deterrent Variables
   PC = (Conditional) Probability of conviction for murder given commission. Defined by $PC = C/Q$, where C = convictions for murder, $Q = M \cdot NS$, NS = state population. This is to correct for the fact that M is an estimate based on a sample from each state.
   PX = (Conditional) Probability of execution given conviction (average number of executions 1946–50 divided by C).
   T = Median time served in months for murder by prisoners released in 1951.
   $XPOS$ = A dummy equal to 1 if $PX > 0$.
c. Independent Economic Variables
   W = Median income of families in 1949.
   X = Percent of families in 1949 with less than one-half W.
   U = Unemployment rate.
   LF = Labor force participation rate.
d. Independent Social and Environmental Variables
   NW = Percent nonwhite.
   AGE = Percent 15–24 years old.
   URB = Percent urban.
   $MALE$ = Percent male.
   $FAMHO$ = Percent of families that are husband and wife both present families.
   $SOUTH$ = A dummy equal to 1 for southern states (Alabama, Arkansas, Delaware, Florida, Kentucky, Louisiana, Maryland, Mississippi, North Carolina, Oklahoma, South Carolina, Tennessee, Texas, Virginia, West Virginia).
e. Weighting Variable
   $SQRTNF$ = Square root of the population of the FBI-reporting region. Note that weighting is done by multiplying variables by $SQRTNF$.
f. Level of Observation
   Observations are for 44 states, 35 executing and 9 nonexecuting. The executing states are: Alabama, Arizona, Arkansas, California, Colorado, Connecticut, Delaware, Florida, Illinois, Indiana, Kansas, Kentucky, Louisiana, Maryland, Massachusetts, Mississippi, Missouri, Nebraska, Nevada, New Jersey, New Mexico, New York, North Carolina, Ohio, Oklahoma, Oregon, Pennsylvania, South Carolina, South Dakota, Tennessee, Texas, Virginia, Washington, West Virginia.
   The nonexecuting states are: Idaho, Maine, Minnesota, Montana, New Hampshire, Rhode Island, Utah, Wisconsin, Wyoming.

[4] This material is taken from a study by a student of mine, Walter McManus (1982).

TABLE 2—ALTERNATIVE PRIOR SPECIFICATIONS

Prior	*PC*	*PX*	*T*	*XPOS*	*W*	*X*	*U*	*LF*	*NW*	*AGE*	*URB*	*MALE*	*FAMHO*	*SOUTH*
Right Winger	*I*	*I*	*I*	*	*D*	*D*	*D*	*D*	*D*	*D*	*D*	*D*	*D*	*D*
Rational Maximizer	*I*	*I*	*I*	*	*I*	*I*	*I*	*I*	*D*	*D*	*D*	*D*	*D*	*D*
Eye-for-an-Eye	*I*	*I*	*D*	*	*D*	*D*	*D*	*D*	*D*	*D*	*D*	*D*	*D*	*D*
Bleeding Heart	*D*	*D*	*D*	*	*I*	*I*	*I*	*I*	*D*	*D*	*D*	*D*	*D*	*D*
Crime of Passion	*D*	*D*	*D*	*	*I*	*I*	*I*	*I*	*I*	*I*	*I*	*I*	*I*	*I*

Notes: 1) *I* indicates variables considered important by a researcher with the respective prior. Thus, every model considered by the researcher will include these variables. *D* indicates variables considered doubtful by the researcher. * indicates *XPOS*, the dummy equal to 1 for executing states. Each prior was pooled with the data two ways: one with *XPOS* treated as important, and one with it as doubtful.

2) With five basic priors and *XPOS* treated as doubtful or important by each, we get ten alternative prior specifications.

the punishment variables to have an effect, but treats all other variables as doubtful. He wants to know whether the data still favor the large deterrent effect, if he omits some of these doubtful variables. The rational maximizer takes the variables that measure the expected economic return of crime as important, but treats the taste variables as doubtful. The eye-for-an-eye prior treats all variables as doubtful except the probability of execution. An individual with the bleeding heart prior sees murder as the result of economic impoverishment. Finally, if murder is thought to be a crime of passion then the punishment variables are doubtful.

In Table 3, I have listed the extreme estimates that could be found by each of these groups of researchers. The right-winger minimum of -22.56 means that a regression of the murder rate data on the three punishment variables and a suitably selected linear combination of the other variables yields an estimate of the deterrent effect equal to 22.56 lives per execution. It is possible also to find an estimate of $-.86$. Anything between these two extremes can be similarly obtained; but no estimate outside this interval can be generated no matter how the doubtful variables are manipulated (linearly). Thus the right winger can report that the inference from this data set that executions deter murders is not fragile. The rational maximizer similarly finds that conclusion insensitive to choice of model, but the other three priors allow execution actually to encourage murder, possibly by a brutalizing effect on society.

TABLE 3—EXTREME ESTIMATES OF THE EFFECT OF EXECUTIONS ON MURDERS

Prior	Minimum Estimate	Maximum Estimate
Right Winger	−22.56	−.86
Rational Maximizer	−15.91	−10.24
Eye-for-an-Eye	−28.66	1.91
Bleeding Heart	−25.59	12.37
Crime of Passion	−17.32	4.10

Note: Least squares is -13.22 with a standard error of 7.2.

I come away from a study of Table 3 with the feeling that any inference from these data about the deterrent effect of capital punishment is too fragile to be believed. It is possible credibly to narrow the set of assumptions, but I do not think that a credibly large set of alternative assumptions will lead to a sharp set of estimates. In another paper (1982), I found a narrower set of priors still leads to inconclusive inferences. And I have ignored the important simultaneity issue (the death penalty may have been imposed in crime ridden states to deter murder) which is often a source of great inferential fragility.

VI. Conclusions

After three decades of churning out estimates, the econometrics club finds itself under critical scrutiny and faces incredulity as never before. Fischer Black writes of "The Trouble with Econometric Models." David

Hendry queries "Econometrics: Alchemy or Science?" John W. Pratt and Robert Schlaifer question our understanding of "The Nature and Discovery of Structure." And Christopher Sims suggests blending "Macroeconomics and Reality."

It is apparent that I too am troubled by the fumes which leak from our computing centers. I believe serious attention to two words would sweeten the atmosphere of econometric discourse. These are whimsy and fragility. In order to draw inferences from data as described by econometric texts, it is necessary to make whimsical assumptions. The professional audience consequently and properly withholds belief until an inference is shown to be adequately insensitive to the choice of assumptions. The haphazard way we individually and collectively study the fragility of inferences leaves most of us unconvinced that any inference is believable. If we are to make effective use of our scarce data resource, it is therefore important that we study fragility in a much more systematic way. If it turns out that almost all inferences from economic data are fragile, I suppose we shall have to revert to our old methods lest we lose our customers in government, business, and on the boardwalk at Atlantic City.

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A Linear Inverse Demand System

Giancarlo Moschini and Anuradha Vissa

We present an inverse demand system that can be estimated in a linear form. The model is derived from a specification of the distance function which is parametrically similar to the cost function underlying the Almost Ideal Demand System. Simulation results suggest that this linear inverse demand system has good approximation properties.

Key words: Almost Ideal Demand System, demand analysis, distance function, duality.

Introduction

The Almost Ideal Demand System (ALIDS) of Deaton and Muellbauer is one of the most commonly used in applied demand analysis. While the ideal connotation of this model stems from its aggregation properties, it is arguable that one of the main reasons for its popularity is the availability of an approximate version of this system that is linear in the parameters; in fact, it is this linear version of the ALIDS model that is typically estimated (Heien and Wessells; Gould, Cox, and Perali; Moschini and Meilke). The purpose of this article is to illustrate how a linear system for inverse demand equations that resembles the ALIDS model can be derived, and we term this system the Linear Inverse Demand System (LIDS).[1]

Inverse demand functions, where prices are functions of quantities, provide an alternative and fully dual approach to the standard analysis of consumer demand (Anderson), and may be more appropriate when quantities are exogenously given and it is the price that must adjust to clear the market (Barten and Bettendorf). This situation is likely to be of relevance to modeling agricultural demand using data based on frequent time series observations (say monthly or quarterly). The chief advantage of using LIDS to model inverse demands is linearity, which may be useful for some applications (say large demand systems or systems involving dynamic adjustment). Although the parametric structure of the model that we present is similar to that of ALIDS, it does not claim the same aggregation properties. Nonetheless, its simplicity and its approximation abilities, documented in this article, are likely to make the LIDS model suitable for empirical studies.

Duality and the Linear Inverse Demand System

Commonly used demand systems typically are derived from parameterizations of dual representations of preferences through the derivative properties. This approach ensures integrability of the resulting demand equations by construction. To derive an inverse demand system, one can start either from the direct utility function and exploit Wold's identity (which yields ordinary inverse demands), or start from the distance (transformation) function and exploit Shephard's theorem (which yields compensated inverse

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demand functions) (Weymark). As will be clear in what follows, for our purposes it is better to start with the distance function, an alternative representation of preferences which has proved convenient in related contexts (Deaton).

If $U(q)$ is the direct utility function, where q denotes the vector of quantities, the transformation or distance function $F(u, q)$ is implicitly defined by $U[q/F(u, q)] \equiv u$, where u is the reference utility level. Under standard regularity conditions, $F(u, q)$ is continuous in (u, q), decreasing in u, and nondecreasing, concave, and homogeneous of degree one in q. These properties establish a useful parallel between the distance function and the cost function $C(u, p)$ derived from the utility-constrained expenditure minimization problem (where p is the price vector corresponding to q). As Blackorby, Primont, and Russell put it (p. 27), ". . . except for the direction of monotonicity of the utility variable, these conditions suggest that C could be interpreted as a transformation function and F as a cost function."

The parallel features of cost and distance functions are useful because, as emphasized by Hanoch, they imply that any standard functional form for the cost function can be applied also to the distance function.[2] The preceding discussion is pertinent to the problem at hand because the useful linear form of the approximate ALIDS model is made possible by the specific functional form chosen for the cost function. Exchanging the role of the variables (u, p) in the PIGLOG cost function of the ALIDS model with the variables $(-u, q)$ of the distance function, where the negative sign on u emphasizes the opposite monotonicity direction of F and C relative to the utility index, one obtains the following parametric specification for $F(u, q)$:

$$\ln(F) = a(q) - ub(q), \tag{1}$$

where $a(q)$ and $b(q)$ are quantity aggregator functions defined as:

$$a(q) = \alpha_0 + \sum_i \alpha_i \ln(q_i) + \frac{1}{2} \sum_i \sum_j \gamma_{ij} \ln(q_i) \ln(q_j), \tag{2}$$

$$b(q) = \beta_0 \prod_i q_i^{\beta_i}. \tag{3}$$

Because $F(u, q)$ is homogeneous of degree one in q, the following restrictions apply: $\Sigma_i \alpha_i = 1$, $\Sigma_j \gamma_{ij} = \Sigma_i \gamma_{ij} = 0$, and $\Sigma_i \beta_i = 0$. Also, without loss of generality, $\gamma_{ij} = \gamma_{ji}$ (the symmetry property).

From Shephard's theorem, the first derivatives of the distance function yield compensated inverse demands as $\pi_i = \partial F / \partial q_i \equiv h_i(u, q)$, where $\pi_i \equiv p_i/x$ is the normalized price of the ith good (the nominal price divided by total expenditure x). Because at $F = 1$ the distance function is an implicit form of the direct utility function, then (1) implies the utility function $U(q) = a(q)/b(q)$. This, together with the derivative property, implies that the uncompensated inverse demand functions associated with (1)–(3) can be written in share form as:

$$w_i = \alpha_i + \sum_j \gamma_{ij} \ln(q_j) - \beta_i \ln(Q), \tag{4}$$

where $w_i \equiv \pi_i q_i$ is the ith budget share, and $\ln(Q)$ is a quantity index defined as $\ln(Q) \equiv a(q)$.

The distance function in (1)–(3) has the same parametric structure of the PIGLOG cost function of the ALIDS model. It should be clear, however, that this distance function is not dual to the PIGLOG cost function of the ALIDS model. It follows that the aggregation properties of the ALIDS are not shared by the inverse demand system in (4). Hence, the attribute "Almost Ideal," used by Eales and Unnevehr and by Barten and Bettendorf to label (4), is somewhat misleading and does not appear warranted for this inverse demand model.

Equations (2) and (4) together entail a nonlinear structure for the inverse demand model.

In practice, however, $\ln(Q)$ can be replaced by an index $\ln(Q^*)$ constructed prior to estimation of the share system to yield:

$$w_i = \alpha_i + \sum_j \gamma_{ij}\ln(q_j) - \beta_i\ln(Q^*). \tag{5}$$

The resulting set of equations (5) is a linear system of inverse demands, the LIDS model. Many index formulae for $\ln(Q^*)$ may be considered here. Similar to the original suggestion of Deaton and Muellbauer, one may use the geometric aggregator $\ln(Q^*) = \Sigma_i w_i\ln(q_i)$, although other indices (say Diewert's superlative indices) may have better approximation properties. It should be understood, however, that in general quantities must be properly scaled for the geometric aggregator to be admissible. This point also applies to the equivalent price aggregator of direct ALIDS models, typically referred to as the Stone price index.[3]

The inverse demand system presented here satisfies standard flexibility properties. It can be verified that the distance function (2)–(4) has enough parameters to be a flexible functional form for an arbitrary distance function once it is realized that the ordinality of utility always allows one to put $\partial^2\ln(F)/\partial u^2 = 0$ at a point.[4]

The notion of flexible functional form in demand perhaps is defined more usefully in terms of demand functions (which are ultimately estimated) rather than in terms of the function representing preferences (which are unobservable). Hence, a flexible inverse demand system must have enough parameters to approximate, at a point, an arbitrary set of quantity elasticities and of normalized price levels (i.e., it must provide a first order local approximation to an arbitrary inverse demand system). If n is the number of goods, it is verified that (after imposing homogeneity, adding-up, and symmetry) the demand system (5) has $\frac{1}{2}(n-1)(n+4)$ free parameters [$(n-1)$ parameters α_i, $(n-1)$ parameters β_i, and $\frac{1}{2}n(n-1)$ parameters γ_{ij}]. These constants could be chosen to represent at a point an arbitrary set of quantity elasticities [of which $\frac{1}{2}n(n+1) - 1$ are independent after accounting for homogeneity, adding-up, and symmetry] and an arbitrary set of left-hand-side shares [of which $(n-1)$ are independent after accounting for adding-up].

Simulation Results

To illustrate the approximation properties of the LIDS model, we report the results of a small simulation exercise. Specifically, we generate repeated stochastic realizations from a known structure and then look at how close the elasticity estimates from LIDS are to the true ones. Following similar studies by Kiefer and MacKinnon, and Wales, the data generating model chosen is a Linear Expenditure System (LES). Specifically, shares for a three-good system are generated using the inverse share equations of LES; that is,

$$w_i = \frac{\alpha_i[q_i/(q_i - \gamma_i)]}{\sum_j \alpha_j[q_j/(q_j - \gamma_j)]}, \tag{6}$$

where $\Sigma_i \alpha_i = 1$. The quantity data that we use for q_1, q_2, and q_3 are U.S. per-capita demand of beef, pork, and chicken, respectively, for the period 1960–89. These data, normalized to equal one at the mean of the sample period, are reported in the appendix.[5] The parameters used are: $\alpha_1 = .5$, $\alpha_2 = .3$, $\alpha_3 = .2$, $\gamma_1 = .2$, $\gamma_2 = .3$, and $\gamma_3 = -.3$. From this structure we generated 250 samples, each with 30 observations, by appending multinormal disturbances to the shares. The (full) covariance matrix used to generate the multinormal errors is the same as that used by Kiefer and MacKinnon, and Wales; that is,

$$\begin{bmatrix} .000036 & -.000025 & -.000011 \\ -.000025 & .000049 & -.000024 \\ -.000011 & -.000024 & .000035 \end{bmatrix}. \tag{7}$$

With these data, we estimate five different models 250 times. First, we estimate the nonlinear inverse demand system of equation (4), and we label this system NLIDS. Similar to the case of ALIDS discussed by Deaton and Muellbauer, the parameter α_0 is virtually impossible to estimate, so we set $\alpha_0 = 0$.[6] Second, we estimate the LIDS model, that is equation (5) with the geometric index $\ln(Q^*) = \Sigma_j w_j \ln(q_j)$. Third, as a benchmark, we estimate the true LES model of equation (6). Note that while LES has five free parameters, both NLIDS and LIDS have seven free parameters. Finally, for comparison, we estimate two versions of the inverse translog (ITL) system introduced by Christensen, Jorgenson, and Lau, and applied by Christensen and Manser, which, after an arbitrary normalization of parameters, can be written as:

$$w_i = \frac{\alpha_i + \sum_j \beta_{ij} \ln(q_j)}{1 + \sum_j \sum_i \beta_{ij} \ln(q_j)}, \tag{8}$$

where $\Sigma_i \alpha_i = 1$ and $\beta_{ij} = \beta_{ji}$.

It can be verified that the ITL system has eight parameters, one more parameter than the LIDS model. Hence, ITL has one more parameter than is needed to make it a flexible (local) approximation to an arbitrary utility, which means that (8) could be suitably restricted without affecting its flexibility properties. Specifically, one can always find a monotonic transformation of utility such that $\Sigma_i \Sigma_j \partial^2 U/\partial \ln(q_i)\partial \ln(q_j) = 0$ at a point. To make this argument more explicit, let $\bar{U}(q)$ denote an arbitrary utility function for which, at a point q^0, $\partial \bar{U}/\partial \ln(q_i) = a_i$ and $\partial^2 \bar{U}/\partial \ln(q_i)\partial \ln(q_j) = a_{ij}$. Because $\bar{U}(q)$ is ordinal, one can put $\Sigma_i a_i = 1$ without loss of generality. Now consider the monotonic transformation $U = G(\bar{U}(q))$. Then, at the point q^0, $\partial^2 U/\partial \ln(q_i)\partial \ln(q_j) = (G'' a_i a_j + G' a_{ij})$, where the derivatives G' and G'' are evaluated at $\bar{U}(q^0)$. Hence, at the point q^0, $\Sigma_i \Sigma_j \partial^2 U/\partial \ln(q_i)\partial \ln(q_j) = G'' + G'(\Sigma_i \Sigma_j a_{ij})$. If one chooses the transformation $G(\cdot)$ such that, at the point q^0, $G' = 1$ and $G'' = -(\Sigma_i \Sigma_j a_{ij})$, then at this point $\Sigma_i \Sigma_j \partial^2 U/\partial \ln(q_i)\partial \ln(q_j) = 0$. Because in the translog utility underlying (8), $\beta_{ij} = \partial^2 U/\ln(q_i)\partial \ln(q_j)$, it follows that we can set $\Sigma_i \Sigma_j \beta_{ij} = 0$ and still have a local approximation to an arbitrary utility function.[7] Given that the translog model (8) with the normalization $\Sigma_i \Sigma_j \beta_{ij} = 0$ achieves what Barnett and Lee called the "minimality" property, the resulting model is termed here the minimal inverse translog (MITL). Like LIDS and NLIDS (with $\alpha_0 = 0$), MITL has seven free parameters.

The approximation properties of the models considered are illustrated in terms of "how close" the estimated elasticities are to the true elasticities. We consider uncompensated quantity elasticities (flexibilities) and scale elasticities (in inverse demand analysis the concept of scale effect, discussed by Anderson, plays a role similar to that of the income effect of direct demands). Quantity elasticities are defined as $\epsilon_{ij} \equiv \partial \ln(p_i)/\partial \ln(q_j)$, and scale elasticities are defined as $\epsilon_i \equiv \partial \ln[p_i(\theta q)]/\partial \ln(\theta)$. Quantity elasticities for LES are computed as:

$$\epsilon_{ij} = w_j\left(\frac{\gamma_j}{q_j - \gamma_j}\right) - \delta_{ij}\left(\frac{q_i}{q_i - \gamma_i}\right), \tag{9}$$

whereas for LIDS and NLIDS they are computed as:

$$\epsilon_{ij} = \frac{\gamma_{ij}}{w_i} - \frac{\beta_i}{w_i}\left(\alpha_j + \sum_k \gamma_{jk} \ln(q_k)\right) - \delta_{ij}, \tag{10}$$

and for ITL and MITL they are computed as:

$$\epsilon_{ij} = \frac{\beta_{ij} - w_i\left(\sum_k \beta_{kj}\right)}{\alpha_i + \sum_k \beta_{ik} \ln(q_k)} - \delta_{ij}, \tag{11}$$

where δ_{ij} is the Kronecker delta ($\delta_{ij} = 1$ for $i = j$ and $\delta_{ij} = 0$ otherwise). Scale elasticities are readily computed using (9)–(11) because $\epsilon_i = \Sigma_j \epsilon_{ij}$.

A possible issue, in light of the arguments presented in Green and Alston, is whether (10) is an appropriate formula for LIDS. It is verified easily that under procedures we have followed (that is, scaling the right-hand-side variables, and setting $\alpha_0 = 0$), each parameter of LIDS will approximate the corresponding parameter of NLIDS. Thus, formula (10), which is derived from NLIDS, is appropriate for LIDS as well. An alternative for LIDS, which is consistent with taking $\ln(Q^*)$ as given in estimation, is to use:

$$\epsilon_{ij} = \frac{\gamma_{ij}}{w_i} - \beta_i \frac{w_j}{w_i} - \delta_{ij}. \tag{12}$$

To investigate what may be called the "local" approximation properties of the model, elasticities were computed at the mean point (at which $q_i = 1 \ \forall \ i$), and summary statistics are reported in table 1.[8] The first column of table 1 reports the elasticities, at the mean point, of the true LES model used to generate the data. Then, for each of LES, ITL, MITL, NLIDS, and LIDS we report the mean, computed over the 250 replications, of the estimated elasticities at the mean point, and the root mean square error (RMSE) of each of these estimated elasticities. Also, for each model we report the average RMSE for the 12 elasticities involved.

All models seem to provide a reasonable approximation. As expected, the best results are obtained by estimating the true LES model. The performances of LIDS, NLIDS, and MITL are similar, with an average RMSE roughly double that of the true model. The fact that MITL does better than ITL perhaps may seem surprising. The reason is that the restriction ($\Sigma_i \Sigma_j \beta_{ij} = 0$) is not rejected; when estimating ITL, the empirical distribution of the quantity ($\Sigma_i \Sigma_j \beta_{ij}$) over the 250 replications had a mean of .4 and a standard deviation of 1.7. Maintaining the restriction ($\Sigma_i \Sigma_j \beta_{ij} = 0$) in MITL results in a considerable efficiency gain (the average absolute t-ratio for the five independent β_{ij} in MITL over the 250 replications was about 3, whereas the average absolute t-ratio for the six independent β_{ij}s in ITL was about 1.4).

Table 1 makes it clear that the linear approximation made possible by the use of $\ln(Q^*)$ instead of $\ln(Q)$ is very good, as LIDS and NLIDS produce virtually identical results. In the context of ALIDS for direct demands, it is believed that the use of the Stone index is likely to produce good approximations because prices typically are highly correlated (Deaton and Muellbauer). In our application, however, the data are not very correlated: the coefficient of correlation between q_1 and q_2 is $-.25$, between q_1 and q_3 is .05, and between q_2 and q_3 is .27. Yet the approximation made possible by the use of $\ln(Q^*)$ appears quite good, suggesting that it is robust to the design matrix of the exogenous variables.

Although the results of table 1 are encouraging as to the approximation properties of LIDS, and consistent with the notion that all the models considered (apart from the true model) are capable of providing a local approximation to an arbitrary demand system, the question arises as to "how local" these results are. If the inverse demand system is to be used for forecasting or welfare analysis, one would want to be reassured that the approximation abilities of the model extend to a reasonably wide range of the data. To investigate this issue, we consider what we term the "extended" approximation properties of the models. Specifically, we evaluate true and estimated elasticities at each of the 30 sample points, and for each of the 12 elasticities we compute the mean square error over the resulting 7,500 estimates (30 sample points for 250 replications).

The square roots of such mean square errors, and their average over all 12 elasticities, are reported in table 2.[9] The approximation abilities of MITL, NLIDS, and LIDS hold up very well in this extended analysis, with the average RMSE increasing only by .004 relative to the approximation at the mean point (up 6.6%). For these models the average RMSE is still roughly twice the RMSE of the true LES model. ITL, on the other hand, shows a much larger increase (up .02 or 30%) in the average RMSE relative to the result at the mean. Again, the restriction ($\Sigma_i \Sigma_j \beta_{ij} = 0$) seems very fruitful in terms of improving the efficiency of the translog inverse demand system.

Table 1. Local Approximation Properties

		LES		ITL		MITL		NLIDS		LIDS	
Elasticity	True Value	Mean	RMSE	Mean	RMSE	Mean	RMSE	Mean	RMSE	Mean	RMSE
ϵ_{11}	−1.121	−1.120	.023	−1.123	.029	−1.130	.025	−1.132	.027	−1.132	.027
ϵ_{12}	.152	.150	.029	.152	.035	.150	.031	.148	.031	.147	.031
ϵ_{13}	−.029	−.030	.005	−.028	.009	−.030	.008	−.030	.008	−.030	.008
ϵ_{21}	.129	.129	.025	.133	.037	.135	.037	.134	.038	.133	.039
ϵ_{22}	−1.276	−1.273	.052	−1.276	.064	−1.278	.063	−1.278	.063	−1.278	.063
ϵ_{23}	−.029	−.030	.005	−.028	.015	−.028	.015	−.027	.015	−.028	.015
ϵ_{31}	.129	.129	.025	.130	.115	.151	.093	.161	.099	.166	.104
ϵ_{32}	.152	.150	.029	.150	.111	.167	.101	.175	.105	.176	.107
ϵ_{33}	−.799	−.798	.032	−.805	.043	−.799	.035	−.800	.035	−.799	.035
ϵ_{1}	−.998	−.999	.039	−.999	.055	−1.010	.043	−1.014	.044	−1.015	.045
ϵ_{2}	−1.176	−1.174	.062	−1.171	.078	−1.171	.078	−1.171	.079	−1.172	.080
ϵ_{3}	−.517	−.519	.046	−.525	.214	−.481	.165	−.464	.175	−.456	.184
Avg. RMSE			.031		.067		.058		.060		.062
Avg. log-likelihood[a]		227.73		229.40		228.68		228.66		228.60	

[a] Average maximized log-likelihood computed assuming that each model in turn is the true model.

Table 2. Extended Approximation Properties

Elasticity	LES	ITL	MITL	NLIDS	LIDS
			RMSE		
ϵ_{11}	.024	.045	.029	.030	.030
ϵ_{12}	.030	.043	.036	.036	.036
ϵ_{13}	.005	.013	.009	.009	.009
ϵ_{21}	.026	.051	.040	.041	.042
ϵ_{22}	.053	.086	.071	.069	.069
ϵ_{23}	.005	.017	.015	.016	.016
ϵ_{31}	.026	.140	.096	.101	.106
ϵ_{32}	.030	.132	.104	.107	.110
ϵ_{33}	.033	.062	.043	.048	.049
ϵ_{1}	.041	.069	.050	.051	.052
ϵ_{2}	.063	.104	.087	.087	.089
ϵ_{3}	.047	.277	.169	.178	.186
Avg.	.032	.087	.062	.064	.066

Note: Entries are RMSEs over all 30 sample points.

Conclusion

In this article we have illustrated a linear specification for an inverse demand system. This specification is based on a distance function which has a parametric structure similar to the PIGLOG cost function underlying the ALIDS model commonly used for direct demand models. The approximation properties of the new model were illustrated with a simulation exercise. Of course, although the results presented are useful in terms of ranking the models used relative to the performance of the true model, the actual size of the approximation error (say, the average RMSE) cannot be generalized because it depends, among other things, on the design matrix of right-hand-side variables, on the structure and parameters of the true model, and on the signal-to-noise ratio of the stochastic terms.

The simulation results show that the new (nonlinear) inverse demand system derived from the chosen parametric specification of the distance function performs well relative to the true model, and very similar to that of an (appropriately restricted) inverse translog demand system. Moreover, the linear version of the new inverse demand system, which we have termed LIDS, results in a good approximation to the nonlinear model. The simplicity of LIDS is likely to make it a useful specification for empirical analysis, especially in applications where linearity is appealing (for example, in dynamic demand systems). Because the derivation of LIDS parallels that of ALIDS for direct demand systems, the simulation results reported in this article are somewhat more general and could be interpreted, with minor modifications, as evidence of the approximation properties of ALIDS models, and as supporting the linear version of ALIDS as a good approximation to the nonlinear ALIDS.

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Notes

[1] After the first draft of this article was completed, a paper by Eales and Unnevehr, giving a similar derivation of the linear inverse demand system, came to our attention. They call this system the "Inverse Almost Ideal Demand System," and use it to model U.S. quarterly meat demand. Barten and Bettendorf also allude to an "Almost Ideal Inverse Demand System," but they do not provide an explicit derivation.

[2] Hanoch formalizes this parallel further by developing the concept of "symmetric" duality, which in our case requires defining the distance function in terms of $(1/u, q)$ rather than (u, q). In Hanoch's words, this approach allows ". . . 'getting two for the price of one' in the search for useful functional forms" (p. 111).

[3] The Stone index fails what Diewert calls the "commensurability test," which defines a fundamental property of index numbers. This property requires that the index should be invariant to the choice of units of measurement.

It is clear that the Stone index, or equivalently the geometric aggregator $\ln(Q^*) = \Sigma_i w_i \ln(q_i)$, is not invariant to the choice of units of measurement. This problem arises when one uses natural units (i.e., pounds or metric tons). In such a situation, an easy way to get around the problem is to scale prices (or quantities for LIDS) by dividing through by the mean. When one aggregate indices with a common base, such as in Deaton and Muellbauer, the problem clearly does not arise.

[4] A similar argument applies to the ALIDS model (Deaton and Muellbauer, p. 313).

[5] These data are from U.S. Department of Agriculture sources. The sample means were 78.417 lbs./capita (retail cut equivalent) for beef, 60.037 lbs./capita (retail cut equivalent) for pork, and 44.283 lbs./capita (ready-to-cook weight) for chicken.

[6] Fixing α_0 basically entails a local normalization of the utility function at the point $q_i = 1$ (the mean point in our case).

[7] The direct demand system derived from an indirect translog utility function also has one more parameter than the linear ALIDS model (or the nonlinear ALIDS with α_0 set to some constant). In this context, imposing the restriction $\Sigma_i \Sigma_j \beta_{ij} = 0$ reduces the indirect translog utility function to be a member of the PIGLOG family of preferences, thereby giving it desirable aggregation properties (Lewbel).

[8] Given that we are evaluating elasticities at the point $q_i = 1$, the distinction between formulae (10) and (12) for LIDS is immaterial, as the two formulae reduce to the same expression at this point.

[9] When evaluating elasticities away from the point $q_i = 1$, formulae (10) and (12) for LIDS are not identical. However, for the three-digit rounding reported in table 2, formulae (10) and (12) give the same results, whereas at a five-digit rounding level formula (10) gives slightly smaller RMSEs.

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Appendix

Table A1. Normalized U.S. Per-Capita Consumption of Beef, Pork, and Chicken

Year	q_1	q_2	q_3
1960	.81870	1.00439	.62778
1961	.83911	.96108	.67520
1962	.84421	.98440	.67294
1963	.89139	1.01605	.69552
1964	.94240	1.01605	.70455
1965	.93858	.91111	.75198
1966	.99596	.90944	.80166
1967	1.01764	.99939	.81972
1968	1.04570	1.02937	.82424
1969	1.05207	1.00938	.86037
1970	1.07630	1.03104	.90553
1971	1.06738	1.13098	.90553
1972	1.09033	1.03936	.93715
1973	1.02657	.94942	.90779
1974	1.08905	1.02271	.91457
1975	1.12221	.84115	.90102
1976	1.20255	.89279	.95973
1977	1.16557	.92943	.98909
1978	1.11201	.92943	1.04780
1979	.99469	1.06102	1.13587
1980	.97428	1.13431	1.12458
1981	.98321	1.08101	1.15845
1982	.97938	.97440	1.19006
1983	.99724	1.03104	1.20587
1984	.99596	1.02437	1.24652
1985	1.00489	1.03270	1.30072
1986	.99979	.97607	1.32556
1987	.93603	.98440	1.41588
1988	.91945	1.05102	1.45653
1989	.87736	1.04270	1.53782

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