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THE ROLE OF EXTERNAL ECONOMIES IN U.S. MANUFACTURING

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ABSTRACT

This paper develops a method for joint estimation of both the degree of internal returns to scale and the extent of external economies. We apply the method in estimating returns to scale indexes for U.S. manufacturing industries at the two-digit level. Overall, we find that only three of the twenty industry categories show any evidence of internal increasing returns: (1) Primary Metals, (2) Electrical Machinery, and (3) Paper Products. More striking, however, is the very strong evidence of the existence of external economies, where external is defined as external to a given two-digit industry and internal to the U.S.. According to our preferred estimates, if all manufacturing industries simultaneously raise their inputs by 10%, aggregate manufacturing production rises by 13%, of which about 5% is due to external economies. Thus, when an industry increases its inputs in isolation by 10%, its output rises by no more than 8%*.

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1. INTRODUCTION

Both in the strict sense and in a figurative sense, external economies have been the driving force behind a considerable body of recent theoretical work.[1] As yet, however, the corresponding empirical literature lags behind. This paper addresses that lag. The central innovation is the development of a method for joint estimation of both the degree of internal returns to scale and the extent of external economies. We apply the method in estimating returns to scale indexes for U.S. manufacturing industries at the two-digit level. Overall, we find very little evidence of internal increasing returns. Of the twenty industry categories only three show any evidence of internal economies. More striking, however, is the very strong evidence of the existence of external economies, where external is defined as external to a given two-digit industry and internal to the U.S.. That is, for a given input level, an industry's output is significantly higher on average when output of other manufacturing industries is high.

The starting point for our modelling approach draws heavily on Hall's ingenious work (1988a, 1988b), in which he estimated indexes of returns to scale, γ, and monopolistic competition, μ, for U.S. industry. His approach, however, is not aimed at capturing the potential role of external economies in production. In fact, if his measured γ is interpreted as the elasticity of industry output with respect to total industry's input then the estimate is upward-biased when external economies are present (given the positive correlation between industry inputs and aggregate inputs[2]). In response to this issue, we develop a model which successfully discriminates between the two types of scale effects. Aggregation is explicitly dealt with in order to determine how external economies become internal as the aggregation level rises. In the end, our point estimates for the degree of internal returns to scale at the two-digit level suggest that increasing returns are both much more muted and much less prevalent than Hall's point estimates, when misinterpreted, suggested. The difference is principally attributable to the general significance of external economies.

In the paper we present a number of different estimating procedures and specifications. With respect to the estimating procedures, before proceeding we explicitly consider the bias and efficiency tradeoffs of instrumental and non-instrumental variable procedures. Although in the end the magnitudes of the estimates do depend on the estimation technique used, the main conclusions do not: The evidence points in the direction of constant or decreasing returns to scale at the two-digit level, with an important external economy that lifts the returns to scale index at the one-digit level above one. According to our preferred estimates, if all manufacturing industries simultaneously

[1] See, for example, Ethier (1982), Romer (1986), and Lucas (1988).

[2] In Hall's case, the source of upward bias is the correlation between aggregate demand instruments and external economies.

raise their inputs by 10%, aggregate manufacturing value added rises by 13%, of which about 5% is due to external economies. Thus, when an industry increases its inputs in isolation by 10%, its value added rises by no more than 8%.

The theoretical literature on the distinction between internal and external economies dates back to Marshall (1920). The distinction is fundamental since the role of increasing returns cannot be dealt with unless their nature is specified, given their very different implications for firm behavior. In his text Marshall provided two examples of external effects. The first is an increase in "trade knowledge" that cannot be kept secret. The second is based on subsidiary trades that use what he referred to as "machinery of the most highly specialized character." This notion harks back to Adam Smith's emphasis on efficient division of labor and the benefits of increased specialization [see also Young (1928)].

Once the idea of an externality was formalized by Meade (1952) and the existence of a consistent, competitive equilibrium with externalities demonstrated by Chipman (1970), it was clear that Marshall's concept could be rigorously justified. It is now widely accepted that, for example, the inclusion of knowledge as an input to production can lead to increasing returns and that spillovers of knowledge between firms can be treated as externalities because protection of proprietary information is incomplete [see, e.g., Arrow (1962) and Romer (1986)[3]][4].

For many years, Marshall's concept of increasing returns that are external to a firm but internal to an industry, or the economy as a whole, was most widely used in static models, especially in the field of international trade [see Helpman (1984) for a discussion of the use of this framework in trade models].[5] The sources for the external effects in these models include advantages of within industry specialization, conglomeration, indivisibilities, and public intermediate inputs such as roads. More relevant for our work than the external to the firm/internal to the industry distinction, however, is the theoretical work that considers inter-industry externalities, since our unit of empirical analysis is the two-digit level industry. Recent work in this area includes Manning and Macmillan (1979), Chang (1981), and Herberg, Kemp, and Tawada (1982). The latter article, for example, considers the effect of (unspecified) inter-industry externalities on the traditional Stolper-Samuelson and

[3] Also see Lucas (1988) for a model with constant returns to scale at the firm level and externalities due to human capital accumulation.

[4] Our procedure is designed to detect technological externalities, as opposed to the pecuniary externalities discussed, for example, in Murphy, Shleifer and Vishny (1988). Note, however, that most of their arguments could be rephrased in terms of technological externalities.

[5] The area of Urban Economics has also frequently made use of Marshall's notion of external economies. For an example of an application of "agglomeration externalities" in a model of the formation of cities see Papageorgiou and Smith (1983).

Rybczynski theorems, among other issues. The authors find that the conclusions from these theorems can only be derived under certain conditions, which typically include non-positive inter-industrial effects.

Of particular interest to Macroeconomists is the fact that the presence of (positive) external economies imparts distinctive characteristics to an economy. For example, business cycle fluctuations are likely to be wider, as co-movements generate important reinforcing cross effects. Furthermore, it is now well known that external effects with positive feedback provides fertile ground for multiple equilibria [see Diamond (1982)].

The rest of the paper is divided into five sections and two appendices. Section II lays out the method in detail and outlines our approach for discriminating between internal and external economies; section III describes the data; section IV presents the analysis of potential bias from not using instrumental variables and the specification test results; section V presents the results for the indexes of returns to scale and external economies; and section VI provides our conclusions. Appendices A and B present extensions and discuss their effects on our parameter estimates.

II. METHOD

II.1. The Basic Model

Consider the following value added function which treats both external economies and technological progress explicitly:

$$Y = F(K, L, E, V),$$

where Y, K, L, E and V are value added, capital, labor, an external economy index and a productivity index, respectively. Furthermore, assume F is homogeneous of degree $\gamma(t)$ in capital and labor, of degree one in the productivity index, and of degree one in the external effects index.[6]

Letting $x = \log X$ and $F_X = \frac{\partial F}{\partial X}$, we get the total differential:

$$dy = \left(\frac{F_K K}{Y}\right) dk + \left(\frac{F_L L}{Y}\right) dl + \left(\frac{F_E E}{Y}\right) de + \left(\frac{F_V V}{Y}\right) dv. \tag{1}$$

Given the homogeneity properties of $F(.,.,.)$, we have:

$$\gamma(t) = \left(\frac{F_K K}{Y}\right) + \left(\frac{F_L L}{Y}\right),$$

[6] This is just a normalization. Since E and V are simply indexes, homogeneity of degree one imposes no constraint.

and

$$1 = \left(\frac{F_E E}{Y}\right) = \left(\frac{F_V V}{Y}\right).$$

Replacing these conditions in (1) yields:

$$dy = \gamma(t)dk + \left(\frac{F_L L}{Y}\right)(dl - dk) + de + dv. \tag{2}$$

Under the very strong assumption that conditions faced by firms are such that their dynamic optimization problem can be well approximated by a sequence of frictionless static problems, it is possible to obtain a simple expression for $(\frac{F_L L}{Y})$, since the first order condition with respect to labor of a profit maximizing firm with some degree of monopoly power in the goods but not in the factor markets is:

$$P\left(\frac{\eta(t) - 1}{\eta(t)}\right) F_L = W,$$

where P, W and η are the value added price, wage and elasticity of demand (absolute value), respectively.

A firm with monopoly power will set its price such that $P/MC = \left(\frac{\eta(t)}{\eta(t)-1}\right)$, where MC denotes marginal cost. Defining the markup coefficient as $\mu(t) \equiv P/MC$ yields:

$$\left(\frac{F_L L}{Y}\right) = \mu(t)\alpha_v(t), \tag{3}$$

where $\alpha_v(t) \equiv \frac{WL}{PY}$ is the share of labor in value added.

Replacing (3) in (2) results in:

$$dy = \gamma(t)dk + \mu(t)\alpha_v(t)(dl - dk) + de + dv. \tag{4}$$

This equation establishes the percent change in output as the sum of four different components. The first is the product of the elasticity of output with respect to total input and the percent change in capital. The second is the elasticity of output with respect to labor, which equals $\gamma(t)\alpha_v(t)$, times the change in the labor/capital ratio. Finally, de is the external effect and dv is the percent change in the productivity index.

Excluding the external economy term, Hall's (1988a) formulation corresponds to the case in which $\gamma(t) = 1$, whereas Hall's (1988b) model is obtained by recognizing that $F(.,.,.)$'s homogeneity of degree $\gamma(t)$ with respect to K and L together with the first order conditions with respect to both factors yields:

$$\frac{PY}{WL + rP_K K} = \frac{\mu(t)}{\gamma(t)}, \tag{5}$$

where r and P_K are the real (rental) cost of capital and price of capital, respectively.

If $\alpha_c(t)$ denotes labor's share in total factor costs, equation (5) implies:

$$\alpha_v(t) = \alpha_c(t)\frac{\gamma(t)}{\mu(t)},$$

hence

$$dy = \gamma\left[\alpha_c(t)dl + (1-\alpha_c(t))dk\right] + dw, \tag{6}$$

where $dw \equiv de + dv$. This is the first estimating equation of this paper (with the coefficient γ assumed time invariant). As discussed above, it matches Hall's (1988b) formulation. It establishes the percent change in output as the weighted percent changes in inputs, multiplied by the returns to scale index γ, plus some non-observable. The weights for the inputs are the corresponding cost shares. Some intuition for the appropriate weights being the cost shares comes from cost minimization. Consider a slight substitution of l for k at the marginal rate of technical substitution (i.e., such that $dy = 0$). Given factor prices, the percent rise in the total labor bill equals dl and the percent fall in the total capital bill equals dk. The only way these (typically) different percent changes can result in no change in total cost is to weight them by their corresponding cost shares. Thus, $[\alpha_c(t)dl + (1-\alpha_c(t))dk] = 0$ when $[F_l dl + F_k dk] = 0$. These two expressions establish a clear link between the cost share and the corresponding marginal product.

The formulation above, however, fails to recognize a potentially very important consideration. The next section shows that when the unobservable includes an external economy component the estimates of γ do not in general represent the degree of homogeneity of the production function with respect to capital and labor, but a combination of this homogeneity and external economies.

II.2. Internal vs. External Economies

II.2.1. General Model

Consider a model in which *total value added* can be divided into a continuum of *sectors* indexed by $j \in [0,1]$, with shares of total value added equal to δ_j. In the same way, sectors are formed by a continuum of *industries* indexed by $i \in [0,1]$, with shares of sector j's value added equal to δ_{ij}. Finally, each industry is divided into a continuum of firms indexed by $h \in [0,1]$, with shares of industry ij's value added equal to δ_{hij}. Adopting the notation:

$$dx_{hij} \equiv \left[\alpha_{c_{hij}} dl_{hij} + (1 - \alpha_{c_{hij}}) dk_{hij}\right],$$

it is possible to rewrite equation (6) for each firm hij as follows:

$$dy_{hij} = \gamma dx_{hij} + de_{hij} + dv_{hij}, \tag{7}$$

where for convenience the time index has been dropped and γ has, for the time being, been assumed equal across firms.

Productivity change can be decomposed into orthogonal components, dv, dv_j, dv_{ij} and du_{1hij}, hence:

$$dv_{hij} = dv + dv_j + dv_{ij} + du_{1hij}. \tag{8}$$

For simplicity external economies can be summarized by linear terms related to higher levels of output aggregation, so:

$$de_{hij} = \beta_{1hij} dy_{ij} + \beta_{2hij} dy_j + \beta_{3hij} dy + du_{2hij}, \tag{9}$$

where constants have been suppressed for convenience and du_{2hij} represents any departure from a deterministic relationship between aggregate value added and external economies[7].

Notice that the "externality" parameters (the β's) can take either positive or negative values. The former corresponds to external *economies* whereas the latter corresponds to external *diseconomies*. For simplicity, these parameters will be assumed to be equal across firms, hence $\beta_{1hij} = \beta_{1h'i'j'}$, $\beta_{2hij} = \beta_{2h'i'j'}$ and $\beta_{3hij} = \beta_{3h'i'j'}$ for all $h, i, j, h', i', j' \in [0,1]$. Appendix A relaxes this assumption.

[7] E.g. a random coefficient model with $\beta_{hij} = \beta + \omega_{hij}$.

Replacing (8) and (9) back in (7) yields:

$$dy_{hij} = \gamma dx_{hij} + \beta_1 dy_{ij} + \beta_2 dy_j + \beta_3 dy + dv + dv_j + dv_{ij} + du_{hij}, \tag{10}$$

where $du_{hij} \equiv du_{1hij} + du_{2hij}$. Now, multiplying both sides of (10) by δ_{hij}, assuming that the production shares are equal to the input shares, and integrating over the support of h, yields industry ij's estimating equation:

$$dy_{ij} = \frac{\gamma}{1-\beta_1} dx_{ij} + \frac{\beta_2}{1-\beta_1} dy_j + \frac{\beta_3}{1-\beta_1} dy + \frac{1}{1-\beta_1}(dv + dv_j + dv_{ij}).$$

The same can be done for higher levels of aggregation, yielding:

$$dy_j = \frac{\gamma}{1-\beta_1-\beta_2} dx_j + \frac{\beta_3}{1-\beta_1-\beta_2} dy + \frac{1}{1-\beta_1-\beta_2}(dv + dv_j)$$

and

$$dy = \frac{\gamma}{1-\beta_1-\beta_2-\beta_3} dx + \frac{1}{1-\beta_1-\beta_2-\beta_3} dv.$$

Certainly extending this to n levels of aggregation is trivial. The formulas reveal how economies that are external at one level become internal at higher levels of aggregation. It also reveals that aggregation does not cause estimation problems as long as the returns to scale estimate is not interpreted as the firm level returns to scale. Nevertheless, estimating disaggregated equations disregarding external economies may induce serious biases if the latter are present. The next subsection develops these ideas in more detail and brings the degree of aggregation and source of external economies to a level at which manufacturing forms a complete system by itself.

II.2.2. Manufacturing Model

In this subsection we lay out the main model to be estimated and tested in this paper. We consider total manufacturing value added as the highest level of aggregation and the two-digit SIC classification as the lowest level of aggregation. In terms of the model presented above, this amounts to assuming that β_3, which captures any effect external to manufacturing as a whole, is small.[8] In addition, the coefficient γ estimated here corresponds to the degree of economies of scale internal to each two-digit industry, and therefore does not necessarily coincide with the degree of economies of scale at the firm level given that economies external to the firm but internal to the industry might be present.

Indexing by i the two-digit industries allows us to write a simplified model with:

$$dy_i = \gamma dx_i + \beta dy + dv + du_i \tag{11}$$

and

$$dy = \frac{\gamma}{1-\beta}dx + \frac{1}{1-\beta}dv, \tag{12}$$

where dy represents value added in manufacturing.

As pointed out above, Hall's estimate [i.e., equation (6)] applied to aggregate manufacturing reflects more than just the average degree of internal returns to scale at the industry level. Given the degree of homogeneity with respect to capital and labor of technologies at the industry level, as the degree of external economies (within manufacturing) rises, the OLS[9] estimate of γ in (6) also rises, reflecting the fact that industry level external effects as specified in our model are internal to the manufacturing sector as a whole.

On the other hand, when equation (6) is estimated at the industry level, the OLS[10] estimate

[8] We did test specifications that admit a potential role for effects external to manufacturing as a whole. That is, we projected industry output changes on GNP changes, in addition to own input changes (dx_i) and total manufacturing output changes (dy). The resulting coefficient on GNP changes was typically insignificant. In those cases in which it was significant the overall fit of the equation was not improved.

[9] Issues of potential correlation between inputs and technological progress are considered in the next subsection. However, the main arguments also hold for the standard instrumental variables procedures used in the literature.

[10] Again, potential correlation between inputs and technological progress are postponed until the next subsection.

of γ, henceforth denoted θ, has a probability limit that in general does not coincide with γ:

$$plim\,\theta = \gamma + \psi\frac{\gamma\beta}{1-\beta}, \tag{13}$$

where $\psi = \frac{\sigma_{dx dx_i}}{\sigma^2_{dx_i}}$.

If, on the other hand, an IV procedure is used:

$$plim\,\theta = \gamma + \psi_{iv}\frac{\gamma\beta}{1-\beta}, \tag{13'}$$

where $\psi_{iv} = \frac{\sigma_{dx d\hat{x}_i}}{\sigma^2_{d\hat{x}_i}}$, and $d\hat{x}_i$ denotes the projection of dx_i onto the subspace spanned by the instruments.

Hence, if industry and total manufacturing inputs are positively correlated ($\psi > 0$) and positive external economies are present ($\beta > 0$), as the results below strongly suggest, θ is an (asymptotically) upward biased estimator of γ. In fact θ reflects an interesting concept: If conditional expectations are assumed to be linear in inputs, θ is a consistent estimator of the expected percent increase in a given industry's output each time this industry increases its inputs by one percent. This expectation, however, takes into account the likelihood of simultaneous input use increases on the rest of the industries. A similar concept applies for the IV estimator.

In contrast, estimation of equation (11) permits a disentangling of internal and external economies. Given the apparent correlation between dy and dv this requires using instrumental variables. Fortunately, equation (12) yields the required instrument for dy.[11] The next subsection describes the estimation procedures.

[11] Do not confuse an instrument for dy with the previous discussion about IV procedures. There, the question was whether to instrument the dx_i's or not.

II.3. Testing Procedure

Estimation of equations (11) and (12) has the inconvenience that (unobservable) productivity growth is likely to be correlated with changes in capital and labor, yielding a classical case of specification error. Hall noticed this and advocated using an instrumental variable procedure. Although theoretically correct, the lack of good macro-instruments rendered his insightful procedure powerless. The reason to worry about specification error is the inconsistency of parameter estimates. However the magnitude of this asymptotic bias decreases with the size of the variance of the regressors relative to their covariance with changes in productivity growth. If the latter is small relative to the former, there is no need to forgo the relative power of OLS or SUR procedures (i.e., relative to IV with instruments mildly correlated with regressors). Furthermore, the same reasoning shows that a very small correlation between instruments and changes in productivity growth may prove much more problematic than the OLS or SUR biases since the covariance between instruments and regressors is likely to be far smaller than the variance of the latter[11]. Section IV presents Hausman's specification tests and provides an example to illustrate the tradeoffs involved.

A different issue arises in estimating equation (12) even when inputs and technological progress innovations are "nearly" independent, since dy and dv are obviously correlated. A 3SLS procedure in which all the x_i's are used as instruments is feasible; however, the shortage of observations leaves very few degrees of freedom in the first stage regression. A more appealing procedure[12] is to use the restricted reduced form in the first stage. In fact the latter is given by equation (12). An equivalent procedure is to just replace equation (12) in (11), yielding:

$$dy_i = \gamma dx_i + \frac{\beta\gamma}{1-\beta}dx + \frac{1}{1-\beta}dv + du_i$$

or

$$dy_i = \gamma dx_i + \kappa dx + \frac{1}{1-\beta}dv + du_i.. \tag{14}$$

The parameter κ represents the external economy in terms of aggregate inputs. Given estimate of κ and γ it is easy to recover the parameter β. Also notice that if manufacturing inputs rather than manufacturing value added is the appropriate externality index (e.g. Romer 1988), so that $de_i = \lambda dx + du_{2i}$, κ captures the magnitude of this external economy as well. In fact our model cannot tell these two stories apart.

[11] Notice that these are all large sample arguments. See Nelson and Startz (1988a, 1988b) for very compelling reasons in the small sample context to prefer OLS procedures over IV with poor first stage regression properties.

[12] Obviously, this is only a small sample issue.

III. DATA

The data used are the same as those used in Hall (1988a, 1988b) and were obtained from Hall. They cover seven one-digit industry groups and the twenty two-digit manufacturing industries for the years 1953-1980. The series we use include: Y: real value added in 1982 dollars [*U.S. National Income and Product Accounts* (NIPA)], K: net real capital stock [Bureau of Economic Analysis], N: hours of work of all employees [NIPA], and W: total compensation divided by N.

The instrumental variables are (1) the rate of increase of the world dollar price of crude petroleum, (2) the rate of growth of military purchases of goods and services in real terms, and (3) a dummy variable with the value of one when the president is a Democrat and zero when he is a Republican.

We construct the rental price of capital as Hall did, followingHall and Jorgenson (1967). The rental price is determined as:

$$r = (\rho + \delta)\frac{1 - c - \tau d}{1 - \tau}p_K$$

where ρ is the firm's real cost of funds, measured as the dividend yield of the S&P 500; δ is the economic rate of depreciation, set to 0.127 [Jorgenson and Sullivan (1981), Table 1, p. 179]; c is the effective rate of the investment tax credit [Ibid., Table 10, p. 194]; d is the present discounted value of tax deductions for depreciation [Ibid., Table 6, pp. 188-189]; and p_K is the deflator for business fixed investment [NIPA].

The standard argument in support of using the dividend yield as the real cost of funds starts with the observation that most investment in the U.S. is financed through equity in the form of retained earnings. As long as the dividend yield is a good measure of the cost of equity, its use is justified. Of course, stocks differ substantially in the share of their total yield that comes from dividends versus capital gains. The argument is that on balance the dividend yield is an accurate measure. The principal alternative is to derive an estimated real rate from some measure of expected or realized inflation and a nominal rate of interest. As it turns out, in Caballero and Lyons (1989) we found that the results in this context are quite robust to the choice of measures.

IV. BIAS ANALYSIS AND SPECIFICATION TESTS

Rightly so, Hall (1988a, 1988b) argues that the returns to scale parameter, γ, is not identified since the changes in inputs are not predetermined with respect to shocks that shift the production function. Unfortunately, instruments that are likely to be uncorrelated with technology shocks are seldom highly correlated with the right hand side variables. Nelson and Startz (1988a, 1988b) showed that when this poor first-stage regression property is accompanied by a small sample, as is certainly true in our case, instrumental variable procedures are potentially more biased than simpler non-instrumental variable approaches.

Unfortunately, these are small sample issues and as such do not have a clear theoretical resolution[13]. Should one use the asymptotically correct but poorly behaved (in small samples) IV estimator[14], or the inconsistent but better behaved (in small samples) non-IV procedures? Given that this is an open question we present both sets of estimates. We show below, nonetheless, that non-IV procedures are probably not subject to substantial asymptotic biases. Furthermore, the large gains in the precision with which the parameters of primary concern are estimated make us lean towards non-IV methods. These are the ones we emphasize throughout the paper.

We first use an example to illustrate that the cost of disregarding the potential correlation between productivity growth and the explanatory variables in terms of asymptotic bias is indeed likely to be small. For this example, let us specialize $F(.,.,.,.)$ to a Cobb-Douglas technology. Let us also assume that the demand conditions are such that the mark-up coefficient, μ, is constant[15] and equal to 1.6. Wages, capital costs, demand and productivity growth are assumed to be driven by independent random variables[16]. Both demand and productivity shocks have idiosyncratic (industry-specific) and aggregate components. Assuming that labor's cost share is 0.75, and taking the conservative approach that all shocks but those to wages and capital costs have the same degree of uncertainty (wages and capital costs are assumed to have a standard deviation twice as large as that of other shocks), it is possible to construct "reasonable" bounds for the potential bias of OLS estimates (i.e., when the possible correlation between regressors and productivity growth is disregarded).

[13] At least there is no clear resolution known to us.

[14] By IV estimator we mean the estimates generated by an instrumental variables procedure in which the instruments are poorly correlated with the regressors.

[15] E.g. an isoelastic demand function.

[16] Some of the independence assumptions are more questionable than others. For example factor price uncertainty is likely to be positively correlated with technology and demand shocks, especially if they are not industry-specific. Under the parametric assumptions made here, however, these assumptions can be relaxed (within "reasonable" margins) without altering results by much.

$$plim(\hat{\phi} - \phi) = \frac{\sigma_{xv}}{\sigma_{xx}} \frac{1}{1-\beta},$$

$$plim(\hat{\gamma} - \gamma) = \frac{1}{[\sigma_{x_i x_i}\sigma_{xx} - \sigma^2_{x_i x}]} \left[\sigma_{xx}\sigma_{x_i v_i} + \left(\frac{\sigma_{x_i x}\sigma_{xv}}{1-\beta}\right)\right],$$

$$plim(\hat{\kappa} - \kappa) = -\frac{1}{[\sigma_{x_i x_i}\sigma_{xx} - \sigma^2_{x_i x}]} \left[\sigma_{xx_i}\sigma_{x_i v_i} - \left(\frac{\sigma_{x_i x_i}\sigma_{xv} - \sigma_{xx_i}\sigma_{x_i v}}{1-\beta}\right)\right]$$

and

$$plim(\hat{\beta} - \beta) = \frac{(1-\beta)plim(\hat{\kappa} - \kappa) - \beta plim(\hat{\gamma} - \gamma)}{(\gamma + \kappa) + plim(\hat{\gamma} - \gamma) + plim(\hat{\kappa} - \kappa)},$$

where

$$\sigma_{xx} = \left[\frac{1-\beta}{\mu(1-\beta) - \gamma}\right] \left[(\mu - 1)^2 + 4(\mu\alpha)^2 + 4(\mu(1-\alpha))^2 + \frac{1}{(1-\beta)^2}\right],$$

$$\sigma_{x_i x_i} = \frac{1 + 4(\mu - 1)^2}{(\mu - \gamma)^2} + \frac{(\mu(1-\beta) + \beta(1-\beta) - \gamma)^2}{(1-\beta)(\mu-\gamma)^2(\mu(1-\beta) - \gamma)}\sigma_{xx},$$

$$\sigma_{xx_i} = \frac{(\mu(1-\beta) + \beta(1-\beta) - \gamma)}{(\mu - \gamma)(\mu(1-\beta) - \gamma)}\sigma_{xx},$$

$$\sigma_{xv} = \frac{1}{\mu(1-\beta) - \gamma},$$

$$\sigma_{x_i v} = \frac{\mu(1-\beta) - \gamma}{(\mu - \gamma)(1-\beta)(\mu(1-\beta) - \gamma)}$$

and

$$\sigma_{x_i v_i} = \frac{1}{\mu - \gamma}.$$

Figures 1a and 1b report the bias involved in OLS estimation for different true values of γ and β. The vertical axes measure the size of the biases and the horizontal axes specify the corresponding value of β, the external economy parameter. The two figures differ only in the true parameter γ assumed. The support of these figures suggests that the biases are never very large[17]. For example,

[17] Especially when compared with the estimated coefficients and standard errors. This holds a fortiori when one considers the standard errors that result when the parameters are estimated with poor macro-instruments.

if $\gamma = 1.0$ and $\beta = 0.3$, the asymptotic bias of OLS estimates are: 0.016 for β, 0.004 for γ, 0.039 for κ and 0.04 for ϕ. All of these are negligible numbers when compared with the true parameter values.

The next step is to provide a statistical metric for the importance of these biases. Hausman's test is particularly suitable for this purpose. This test relies on the comparison of the parameter of primary concern yielded by two alternative procedures leading to consistent estimates under the null hypothesis (no specification error) and diverging under the alternative. Call $\hat{\beta}_1$ and $\hat{\beta}_2$ these estimates, then Hausman's test is:

$$\left(\hat{\beta}_1 - \hat{\beta}_2\right)^T V(\hat{\beta}_1 - \hat{\beta}_2)^+ \left(\hat{\beta}_1 - \hat{\beta}_2\right),$$

where $V(x)^+$ denotes the generalized inverse of the variance-covariance matrix of the vector x, and x^T is the transpose of x.

Under the null hypothesis this statistic is distributed χ^2 with degrees of freedom equal to the rank of the variance-covariance matrix. Most of the complications come from calculating the variance-covariance matrix. Two notable exceptions are the case in which one of the estimators is relatively efficient and the case in which both sets of estimates are uncorrelated. In the former an asymptotic version of Rao-Blackwell's Theorem applies, yielding the variance of the difference equal to the difference of the variances (Hausman 1978). In the latter, the variance of the difference is equal to the sum of the variances (e.g. Hausman and Taylor 1982).

Table 1 presents the χ^2 statistics of Hausman's tests comparing OLS versus SUR estimates, and 3SLS[18] versus SUR estimates[19] for equation (14). The evidence from the tests is mixed: The first case yields a very low value for Hausman's statistic, suggesting no substantial problems due to the potential correlation between changes in productivity growth and the regressors. The second, on the other hand, reflects a very strong rejection of the no-misspecification hypothesis. Once more, the tradeoffs discussed above do not permit a clear answer. Despite the fact that the variance-covariance matrix of cross equation disturbances is nowhere near diagonal, the first

[18] By 3SLS we mean an instrumental variable procedure that takes into account the covariances between cross equation disturbances. The instrumental variables used are (1) the rate of increase of the world price of crude petroleum in dollars, (2) the rate of growth of military purchases of goods and services in real terms, and (3) a dummy variable with the value of one when the president is a Democrat and zero when he is a Republican. The performance of these instruments is further discussed in Hall (1988b).

[19] Both of these sets of tests correspond to the first of the exceptions mentioned above, in which Hausman's test takes a simple form.

test may suffer from low power. And while the second test is likely to be more powerful, it is not immune to criticisms along the lines of those introduced by Nelson and Startz vis-a-vis the small sample properties of the t-statistic when poor instruments are used. It does not seem too farfetched to conjecture that in this context, given the weak correlation between the instruments and regressors, the small sample distribution of Hausman's statistic is quite far from the standard χ^2 distribution. Nevertheless, given these ambiguities, hereafter we report both sets of estimates, though we emphasize the non-IV procedures.

V. EMPIRICAL RESULTS

V.1. Aggregate Manufacturing

Table 2 presents both OLS and IV results for equation (6) at the level of aggregate manufacturing (constant not shown). The sample extends from 1953 to 1980. According to the model, the coefficient represents the elasticity of output with respect to simultaneous proportional changes in K and L. We denote the coefficient as ϕ rather than γ since the presence of external economies calls for the model described by equation (12) in which this coefficient is shown to represent $\gamma/(1-\beta)$, where β captures the external (to the two-digit industries) economies and γ reflects the average degree of internal returns to scale at the industry level.

On the whole the equation fits well under both methods of estimation, although the precision of the coefficient is much higher using OLS[20]. While the IV estimate is larger, only the OLS estimate is significantly greater than one, suggesting the presence of increasing returns to scale at the aggregate manufacturing level. It is not clear from this coefficient, however, to what degree external returns are present. As section II.2. demonstrates, when ϕ is interpreted as a measure of internal returns to scale at the two-digit level it includes an upward bias if external economies are present[21]. To determine whether this is the case we now turn to the two-digit industry level results.

[20] There is no evidence of serial correlation from either the Durbin-Watson statistic or the Ljung-Box Q statistic.

[21] If external diseconomies prevailed the sign of the bias would be reversed.

V.2. Internal vs. External Economies: The Constrained Model

To provide an overview of the results that follow, Table 3 presents estimates of the models with all industry coefficients constrained equal. The table presents the SUR and 3SLS results for both the model described in equation (6) (first row) and the richer model described in equation (14) which discriminates between internal and external effects (rows 2-5). The marginal significance levels of various tests of the constraints appear in rows 6 through 9.

Row 1 of the table provides an individual two-digit industry estimate of the composite measure of returns to scale from equation (6), denoted here (at the industry level) as θ. The coefficients are both very precisely estimated. While the 3SLS estimate is significantly greater than one, the SUR estimate is on the borderline. Economically, however, the SUR estimate is right at one. Given that we view this as the better measure of the two, we find this very encouraging in light of the fact that so much of the literature finds something close to constant returns to scale at the industry level[22]. Row 6 provides the marginal significance of the test that the θ's are in fact equal across all industries in each country. The more powerful SUR test rejects the constraint with a marginal significance at 2 percent. The individual industry θ's are presented in the next section.

Rows 2 through 5 provide estimates relevant for isolating the role of intra-U.S. external effects. Row 2 presents the arithmetic mean of the estimated ψ's and ψ_{iv}'s, two of the parameters that govern the wedge between θ_i and γ_i, where $\psi = \frac{\sigma_{dxdx_i}}{\sigma^2_{dx_i}}$ and $\psi_{iv} = \frac{\sigma_{dxd\hat{x}_i}}{\sigma^2_{d\hat{x}_i}}$. Thus, there exists on average considerable positive covariance between industry input and aggregate input levels[23]. Rows 3 and 4 present the estimates of the internal elasticities of output with respect to K and L and the role of external economies as captured by κ. Row 5 reports the values for β implied by the estimates of κ. Note that the disentangled coefficients are also very sharply estimated. In neither case is γ significantly greater than one. The external economy parameter is positive and very significant. External economies are clearly present. Moreover, the fact that the aggregate estimates of ϕ (Table 2) are significantly larger than the estimates of γ at the industry level, as well as the fact that θ is between these two estimates[24], is fully consistent with equations (11) and (12), confirming further the presence of external economies and the propriety of the model. In sum, the data appear to speak quite clearly as to the extent and nature (i.e., internal vs. external) of returns to scale at the macro level in the United States.

Row 7 presents the marginal significance levels of the tests that the external returns parameters

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[22] See, for example, the survey by Walters (1963).

[23] Both when projected onto the instruments subspace and when not projected.

[24] As it should be according to equation (12) when $\beta > 0$ and $0 < \psi < 1$. In fact $plim\,\theta$ can be written as $plim\,\theta = \phi + (\psi - 1)\gamma\beta/(1-\beta)$.

are equal across all industries. The constraint cannot be rejected at conventional significance levels for either method of estimation. Rows 8 and 9 provide marginal significance levels for tests of the γ constraint across industries. The more powerful SUR test rejects in both cases.

V.3. Internal vs. External Economies: The Industry Estimates

We now turn to individual two-digit industry results. Table 4 presents the coefficients from estimation of equation (14) with the external effects parameter β, hence κ, constrained equal across all sectors[25]. (Recall from Table 3 that this constraint was the only one that was not rejected for either estimation method.) For both procedures the constrained κ, representing the external effects, is positive and very significant. Moreover, the size of the coefficient is quite similar to that in the case where both κ and γ are constrained equal across sectors (Table 3). On the whole, the measured γ's are positive, significant, much more tightly estimated using SUR, slightly higher using 3SLS, and predominantly lower than one. Of the SUR estimates, five of the twenty industries show a coefficient that is not significantly greater than zero, with Food Products as the egregious outlier on the low end. On the high end, across the two estimation methods only three industries show significant evidence of internal increasing returns: (1) Primary Metals, (2) Paper Products, and (3) Electrical Machinery. The SUR procedure has enough power to indicate that 15 of the 20 two-digit industries exhibit decreasing returns to scale. Overall, the model appears to fit the U.S. two-digit data quite well.

The most prominent feature of the above results is the clear statistical significance of the external economies parameter. In considering the economic significance of its magnitude, perhaps it is most informative to calculate the output effect for an industry that expands independently versus expanding in concert with all other (manufacturing) industries. Consider the following example which uses parameter values in the middle range of the estimated coefficients. An industry with a γ equal to 0.8 and a κ equal to 0.5 increases its total inputs by 10 percent. If all industries act in concert then the added output is 13 percent, as opposed to the 8 percent that results if the industry acts alone (assuming this industry is too small to significantly affect the aggregate). Of course, this comparison considers two polar extremes that are in themselves not very realistic. Nonetheless, even for far less extreme comparisons the differential output effects of external economies are not trivial.

Table 5 presents the industry level results from estimation of equation (6). These θ's correspond to the composite returns to scale measure estimated by Hall (1988b). As with the γ's in

[25] Note that the $\gamma/(1-\beta)$ term in κ corresponds to the aggregate coefficient; hence it is the same for all sectors.

Table 4, on the whole they are positive, significant, much more tightly estimated using SUR, and higher using 3SLS. The prominent sectors on the low end and the high end are much the same as those from Table 4. For each industry, the parameters ψ and ψ_{iv} govern the unconditional change in output for a given change in industry input (given the industry's κ): the higher are ψ and ψ_{iv}, the higher the expected output response to a given change in industry input. The mean ψ and ψ_{iv} are 0.79 and 0.68, respectively, as reported in Table 3. Finally, it is important to notice that of the forty returns to scale estimates reported in tables 4 and 5 only one (sector 21's IV estimate) fails to satisfy the sign of the difference $(\theta - \gamma)$ implied by our model. This represents strong support for the external economy hypothesis.

VI. CONCLUSIONS

Our evidence for the presence of external economies stands in sharp relief. Irrespective of the method of estimation, external economies are both statistically and economically very significant. At the same time, we find no evidence of internal increasing returns in over four-fifths of the twenty two-digit industries. In short, the model we develop for disentangling internal and external economies meets the approval of the data. Expressed compactly, our preferred model says that the elasticity of an industry's output with respect to own input, holding the input of other industries constant, is about 0.8. This elasticity rises to 1.3 when industry inputs move in lockstep. Of course, the unconditional elasticity falls somewhere between these two extremes. According to our estimates it is in the neighborhood of 1.0.

We also argue and present evidence as to why the usual instrumental variable procedures, with the attendant dismal macro instruments, may not be the appropriate choice in the context of Hall's (1988b) procedures, given that our main concern is parameter estimation. The results of Hausman tests designed to detect the potential misspecification are mixed. In the end, although we prefer the non-IV estimation, we present results for both methods. More important, the spirit of the results is invariant to which method is used.

It should be noted that there is no sense in which our efforts have been misguided by looking for external effects at the two-digit level. To be sure, if we had found nothing then it would have been time to look to the less aggregated data for external effects that are fully internalized at lower levels. In fact, we view our results as all the more striking in that they highlight the fact that economically quite large external effects are not wholly internalized even at the two-digit level. It is of course possible that there exist effects external to firms but internal to, say, four-digit level industries that are ultimately reflected in our composite γ's at the two-digit level. This is not a problem. One should simply be reminded that it would be improper to interpret our γ's as measures of internal returns to scale at less aggregated levels.

Our level of aggregation also jibes with much of the theoretical work in the area. For example, Romer's (1986) celebrated growth model, in which knowledge spillovers generate external economies, is typically couched in terms of firm level production versus aggregate production. What matters for his results, however, is that the externality lies somewhere between the firm and the aggregate. Thus, our results are fully consistent with his, and might be interpreted as providing some evidence in support of his framework[26].

Finally, a comparison with European industry is possible since Caballero and Lyons (1989) apply the method developed here to European manufacturing at the NACE two-digit level. The bottom line there is very similar to that in the U.S.. The external economies coefficients (κ's) for the four countries covered by the study are all very significant. Two of the four are higher than in the U.S. (France at 1.40 and Belgium at 0.68) and two are lower (Germany at 0.32 and the U.K. at 0.29). Moreover, as in the U.S., there is very little evidence of internal increasing returns in Europe.

[26] Strictly speaking, however, we cannot rule out the possibility of increasing returns at the firm level, since external *diseconomies* may be present at levels between the firm and the two-digit level.

APPENDIX A

This appendix discusses the potential bias involved with the assumptions of equal β's and γ's across sectors. For this let us write a generalized version of equation (11):

$$(A.1) \qquad dy_i = \gamma_i dx_i + \beta_i dy + dv + du_i.$$

As before, multiplying both sides of $(A.1)$ by δ_i, the share of industry i in manufacturing value added (and inputs), and integrating over all i, yields:

$$(A.2) \qquad dy = \frac{1}{1-\bar{\beta}} \int_0^1 \gamma_i dx_i \delta_i \, di + \frac{1}{1-\bar{\beta}} dv,$$

where $\bar{\beta} \equiv \int_0^1 \delta_i \beta_i \, di$.

Therefore even if the β_i's are different across industries, the aggregate has a simple interpretation in terms of the weighted average of the industry level coefficients. Heterogeneous γ_i's, on the other hand, makes the problem more complex.

Define $\bar{\gamma} \equiv \int_0^1 \delta_i \gamma_i \, di$, then $(A.2)$ can be written as follows:

$$(A.3) \qquad dy = \frac{\bar{\gamma}}{1-\bar{\beta}} dx + \frac{1}{1-\bar{\beta}} dv + \frac{1}{1-\bar{\beta}} \int_0^1 (\gamma_i - \bar{\gamma}) dx_i \delta_i \, di.$$

Equation $(A.3)$ allows us to determine the sign of the bias of the aggregate coefficient ϕ, when the γ_i's are assumed equal across sectors. If, on average, industries that have a higher degree of returns to scale have more procyclical input demands than firms with a lower degree of returns to scale, the estimator of ϕ is upward biased. The opposite is true if firms with higher γ's have less procyclical input demands than firms with lower γ's. Finally, no bias occurs if there is no systematic relation between the γ's and the cyclical behavior of the dx_i's.

From $(A.3)$ it is possible to show, after some simple algebra and assuming that the number of sectors is large, that:

$$(A.4) \qquad plim \left(\frac{\hat{\phi} - \phi}{\phi} \right) = \mathrm{E}\left[\left(\frac{\gamma_i - \bar{\gamma}}{\bar{\gamma}} \right) \nu_i \right],$$

where $\nu_i \equiv \psi_i \sigma_{x_i}^2 / \sigma_x^2$. Several experiments can now be performed. Suppose for instance that the actual γ_i's and ν_i's are jointly normally distributed. Estimating the first two moments from our data yields an asymptotic bias of the above proportion equal to 2.6%.

APPENDIX B

In the paper we have assumed that the dynamic optimization problem of a firm can be well approximated by a sequence of (annual) *frictionless* static problems. This appendix shows that relaxing this assumption to allow for issues like labor hoarding and excess capacity is not likely to yield substantial biases.

Suppose now that the first order conditions implied by the static optimization problem are not always satisfied. Furthermore, they need not even be satisfied on average. After some trivial algebraic manipulations this situation can be described by the following relationships:

$$\mathrm{E}\left[\frac{F_L L}{Y}\right] = \lambda^l \mu \alpha_v(t)$$

and

$$\mathrm{E}\left[\frac{F_K K}{Y}\right] = \lambda^k \mu \alpha_v(t),$$

or

$$\frac{F_L L}{Y} = \lambda^l \mu \alpha_v(t) z^l(t)$$

and

$$\frac{F_K K}{Y} = \lambda^k \mu \alpha_v(t) z^k(t),$$

where λ^l and λ^k are two positive constants, and z^l and z^k are strictly positive random variables with mean equal to one.

If, for example, $\lambda^k < 1$ there is chronic excess capacity. The random variable z^k, on the other hand, controls short-run excess capacity and/or over-utilization of capital. Analogous examples can be constructed for labor hoarding.

Following steps identical to those in the paper we obtain an equation parallel to (4):

$$(B.1) \qquad dy = \gamma(t) dk + \lambda^l z^l(t) \mu(t) \alpha_v(t)(dl - dk) + de + dv.$$

However there is also a parallel to equation (5):

$$\frac{PY}{WL + \tau(t) r P_K K} = \frac{\mu(t)}{\gamma(t)} \lambda^l z^l(t),$$

where $\tau(t) \equiv (\lambda^k z^k / \lambda^l z^l)$. Hence,

$$\alpha_v(t) = \left(\frac{\alpha_c(t)}{\tau(t) + \alpha_c(t)(1 - \tau(t))}\right) \frac{\gamma(t)}{\mu(t)} \frac{1}{\lambda^l z^l(t)}.$$

Plugging this expression back in $(B.1)$, yields:

$$(B.2) \qquad dy = \gamma(t)dk + \left(\frac{\alpha_c(t)}{\tau(t) + \alpha_c(t)(1-\tau(t))}\right)(dl - dk) + de + dv.$$

It is apparent from $(B.2)$ that if $\tau = 1$ the estimating equation is identical to that of the paper. In other words, excess capacity and labor hoarding affect the model only if they do not move proportionally. Hence, the case of proportional excess capacity and labor hoarding (positive or negative) movements is fully consistent with the procedure pursued in the paper.

Now assume that this is not the case, i.e. τ is not always equal to one. To make the formulae simple, however, let us assume that τ is constant, but it takes on very extreme values: when excess capacity is the dominant feature $\tau = 0.5$, whereas when labor hoarding dominates $\tau = 2$. Furthermore, given that only the order of magnitude of the bias concerns us here, let us disregard the external effect (alternatively, assume that the firms capital/labor ratio is almost uncorrelated with aggregate inputs). Finally, assume that changes in a firm's capital/labor ratio are independent of the rate of change of its capital, and the latter is twice as volatile as the former[27]. In this case, the asymptotic bias of the OLS estimate of γ is:

$$plim\left(\frac{\hat{\gamma}-\gamma}{\gamma}\right) = \frac{\alpha_c^2(1-\alpha_c)(1-\tau)}{4(1+\alpha_c^2/4)(\tau+\alpha_c(1-\tau))}.$$

This expression takes a value equal to 2.3% for $\tau = 0.5$ and -3.3% for $\tau = 2.0$, very small numbers. Adding uncertainty respect to τ, regardless of whether it is pro- or counter-cyclical, does not change the order of magnitude of these biases.

[27] Given that capital and labor tend to comove, this seems to be a very conservative assumption. In fact the asymptotic bias is increasing in the variance of the capital/labor ratio.

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TABLE 1

Hausman Tests: Chi–Square Statistics for Both Types of Test

$$\text{Eqn (14): } dy_i = \gamma_i dx_i + \kappa_i dx + \left[\frac{1}{1-\beta} dv + du_i\right]$$

	OLS–SUR	SUR–3SLS
$\beta_i=\beta_j$, γ_i Unconstr.	3.2	45.1**
$\beta_i=\beta_j$ & $\gamma_i=\gamma_j$	0.1	24.7**

** : Significant at the 1% level.
* : Significant at the 5% level.

TABLE 2

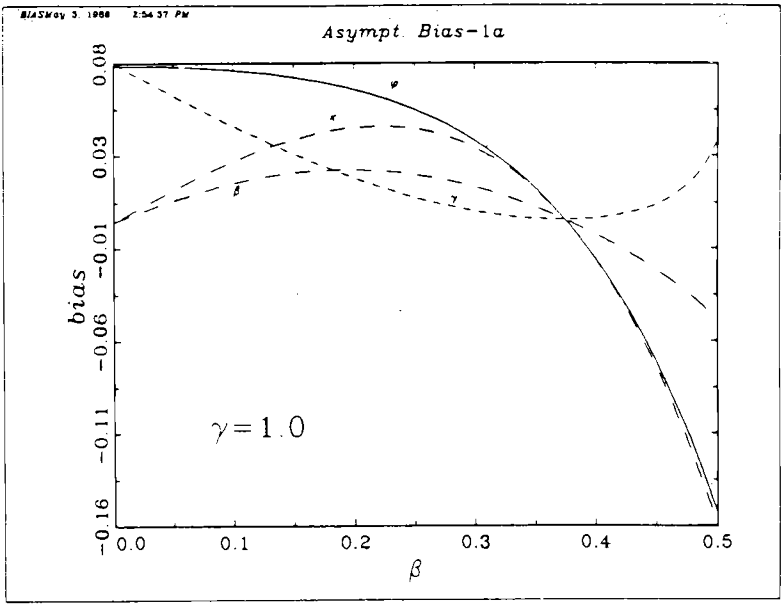
The Extent of Increasing Returns in Aggregate Manufacturing

$$\text{Eqn (6): } dy = \phi[\alpha_c(t)dl + (1-\alpha_c(t))dk] + dw$$

		OLS	IV
Agg. Manufacturing	ϕ	1.37 (0.08)	1.88 (0.46)

*Standard errors in parentheses.

Figures 1a and 1b: Bias Analysis



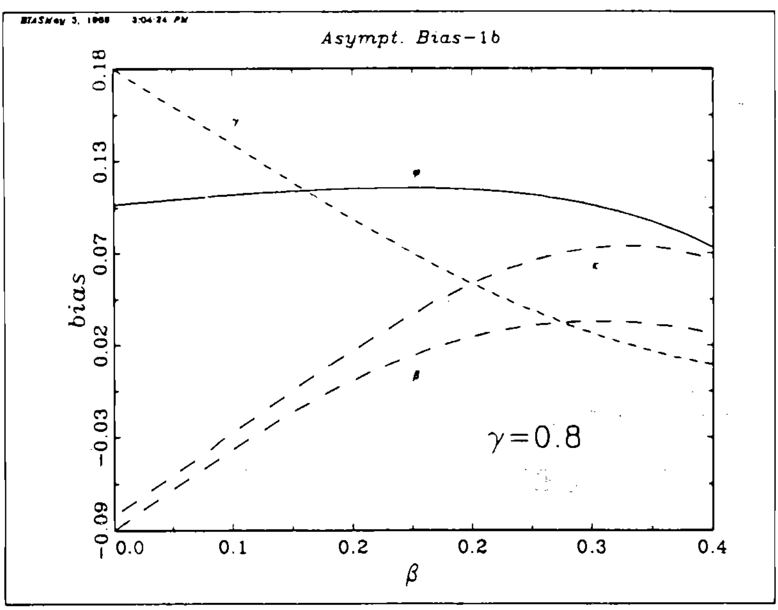


TABLE 3

Internal vs. External Economies: Results for Constrained Model

$$\text{Eqn (6): } dy_i = \theta_i dx_i + [de_i + dv_i]$$

$$\text{Eqn (14): } dy_i = \gamma_i dx_i + \kappa_i dx + [(1/(1-\beta))dv + du_i]$$

$$\text{plim } \theta = \gamma + \psi[\beta\gamma/(1-\beta)]$$

$$\kappa \equiv \beta\gamma/(1-\beta)$$

	SUR	3SLS
θ	1.04 (0.02)	1.32 (0.06)
$\bar{\psi}$	0.79	0.68
γ	0.78 (0.02)	0.93 (0.05)
κ	0.49 (0.05)	0.89 (0.13)
β	0.39	0.49
P–Value $\theta_i = \theta_j$	0.02	0.83
P–Value $\beta_i = \beta_j$	0.07	0.27
P–Value $\beta_i = \beta_j$ & $\gamma_i = \gamma_j$	0.00	0.56
P–Value $\gamma_i = \gamma_j$ given that $\beta_i = \beta_j$	0.01	0.83

*Standard errors in parentheses.

TABLE 4: Internal vs. External Economies at the Two–Digit Level

	γ (SUR)	κ (SUR)	γ (3SLS)	κ (3SLS
Food Products (20)	–1.14 (0.36)	0.44 (0.06)	–1.10 (0.82)	0.71 (0.23)
Tobacco Products (21)	0.02 (0.16)	"	5.24 (2.33)	"
Textile Products (22)	0.49 (0.18)	"	1.34 (0.45)	"
Apparel (23)	0.51 (0.09)	"	0.65 (0.20)	"
Wood Products (24)	0.49 (0.11)	"	0.94 (0.27)	"
Furniture (25)	0.98 (0.06)	"	1.04 (0.14)	"
Paper Products (26)	1.35 (0.18)	"	1.87 (0.49)	"
Printing/Publishing (27)	0.33 (0.17)	"	0.42 (0.40)	"
Chemical Products (28)	0.07 (0.18)	"	0.36 (0.75)	"
Petroleum Refining (29)	–0.32 (0.16)	"	0.02 (0.99)	"
Rubber/Plastic (30)	0.83 (0.08)	"	1.26 (0.18)	"
Leather Products (31)	1.14 (0.13)	"	0.14 (0.34)	"
Stone/Clay/Glass (32)	0.77 (0.09)	"	1.29 (0.32)	"
Primary Metals (33)	1.65 (0.07)	"	1.82 (0.33)	"
Fabricated Metals (34)	0.67 (0.06)	"	0.83 (0.38)	"
Machinery (35)	0.95 (0.05)	"	0.93 (0.16)	"
Electrical Machinery (36)	0.76 (0.07)	"	1.61 (0.27)	"
Transport Equip. (37)	0.78 (0.08)	"	0.90 (0.24)	"
Instruments (38)	0.48 (0.09)	"	0.53 (0.39)	"
Misc. Manufacturing (39)	0.32 (0.13)	"	0.63 (0.73)	"

*Standard errors in parentheses.

TABLE 5: Composite Returns to Scale at the Two—Digit Level

	θ (SUR)	θ (3SLS)	ψ	ψ(IV)
Food Products (20)	0.09	0.50	2.89	2.41
	(0.39)	(0.90)	(0.45)	(0.98)
Tobacco Products (21)	–0.12	4.34	–0.37	0.02
	(0.18)	(2.28)	(0.30)	(1.78)
Textile Products (22)	0.86	2.12	0.73	0.59
	(0.18)	(0.44)	(0.09)	(0.18)
Apparel (23)	0.87	1.02	0.79	0.60
	(0.09)	(0.22)	(0.10)	(0.21)
Wood Products (24)	0.77	1.25	0.43	0.28
	(0.11)	(0.26)	(0.08)	(0.17)
Furniture (25)	1.24	1.32	0.57	0.41
	(0.06)	(0.16)	(0.06)	(0.14)
Paper Products (26)	1.78	2.52	1.33	1.24
	(0.18)	(0.47)	(0.16)	(0.38)
Printing/Publishing (27)	1.07	1.48	1.57	1.50
	(0.14)	(0.39)	(0.25)	(0.55)
Chemical Products (28)	0.53	1.65	1.23	1.05
	(0.19)	(0.88)	(0.28)	(0.68)
Petroleum Refining (29)	–0.25	–1.43	0.13	–1.49
	(0.14)	(1.13)	(0.34)	(1.14)
Rubber/Plastic (30)	0.98	1.59	0.57	0.48
	(0.08)	(0.21)	(0.04)	(0.11)
Leather Products (31)	1.34	0.47	0.67	0.57
	(0.13)	(0.29)	(0.12)	(0.22)
Stone/Clay/Glass (32)	1.11	1.75	0.93	0.54
	(0.10)	(0.29)	(0.10)	(0.29)
Primary Metals (33)	1.91	2.18	0.65	0.88
	(0.08)	(0.35)	(0.06)	(0.26)
Fabricated Metals (34)	0.92	1.64	0.59	1.01
	(0.07)	(0.54)	(0.07)	(0.49)
Machinery (35)	1.15	1.31	0.57	0.53
	(0.04)	(0.13)	(0.05)	(0.16)
Electrical Machinery (36)	0.95	2.14	0.58	0.63
	(0.07)	(0.31)	(0.04)	(0.15)
Transport Equip. (37)	0.90	0.99	0.43	0.35
	(0.10)	(0.26)	(0.08)	(0.15)
Instruments (38)	0.70	1.16	0.60	0.82
	(0.09)	(0.50)	(0.07)	(0.45)
Misc. Manufacturing (39)	0.54	1.64	0.92	1.10
	(0.15)	(0.86)	(0.08)	(0.27)

*Standard errors in parentheses.