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ABSTRACT

The success of deregulation in creating a viable private rail freight system in the U.S. since 1979 is relatively undisputed. Deregulation has proceeded in three ways: (i) eased rate setting restrictions; (ii) simplified merger applications and approval procedures; and (iii) relaxed route abandonment policies.

In this paper we attempt to disentangle the effects of deregulation on rail costs from those directly attributable to mergers and acquisitions. We employ a translog variable cost function, based on an unbalanced panel data set of annual observations for major U.S. Class I railroads from 1974 to 1986.

We find that both deregulation and mergers contributed significantly to cost savings. However, of the accumulated cost savings achieved by the six major firms involved in mergers post-deregulation, we estimate that by 1986 about 91% of the reduction in accumulated costs is due to deregulation, and about 9% is directly due to mergers and acquisitions (which in turn were facilitated by regulatory reforms). In terms of factor biases, we find that both deregulation and mergers resulted in a substantial labor-saving bias; the point estimate of the deregulation labor-saving bias is larger than that for mergers, but we were not able to estimate this bias precisely.

We conclude that mergers were not a prerequisite for railroads being able to achieve substantial cost and productivity improvements in our 1974-1986 sample period. Deregulation also had an enormous direct impact; indeed, its impact appears to have been much larger.

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I. INTRODUCTION

In the decade since the passage of the Staggers' Act -- legislation that substantially deregulated U.S. railroads -- considerable change has occurred. The number of Class I railroads has diminished through bankruptcies, consolidations and mergers; concentration has risen concomitantly, with seven carriers now accounting for over 70% of rail freight traffic. Over 25,000 miles of track have either been sold or abandoned; and rates of return have risen substantially, with several railroads now achieving revenue adequacy.

While the impacts of deregulation on the railroads are widely recognized, the sources of the changes that have occurred in this industry have not been fully analyzed. Because deregulation gave the railroads substantial freedom in rate setting, track abandonments and mergers, it provided them with a number of avenues through which costs could be reduced and revenue could be enhanced.

As part of an on-going research project, we have attempted to disentangle the various effects of deregulation upon the railroads. In particular, the impact of deregulation upon rates and revenues on the one hand, and costs and capital adjustments on the other, have been respectively analyzed by Friedlaender [1990], and by Friedlaender, Berndt, Chiang and Vellturo [1990]. In this paper we attempt to disentangle the effects of deregulation on real costs and productivity from those directly attributable to mergers and acquisitions. Our empirical analysis employs a short-run variable cost function, and our econometric estimates are based on an unbalanced panel data set of annual observations for major Class I U.S. railroads from 1974 to 1986.

This paper takes the following form. Section II discusses the nature of rail mergers, specifies an econometric model based on a translog short-run variable cost function and then outlines how the effects of deregulation and mergers on costs and productivity growth can be distinguished and quantified. Section III describes the data set, while Section IV provides a discussion of a number of econometric issues related to the specification and estimation of the cost function equations. Our principal empirical findings are given in Section V, and in the final section we briefly summarize our conclusions and suggest issues for future research.

II. RAIL MERGERS AND PRODUCTIVITY GROWTH

Prior to the onset of deregulation, rail mergers typically involved unions of systems having a great deal of parallel trackage.[1] Under such circumstances, hearings at the Interstate Commerce Commission (ICC) typically focused on the traditional welfare trade-off that mergers could provide between the cost efficiencies occasioned by consolidation and economies of scale on the one hand, and the potential for noncompetitive pricing on the other. These hearings were often lengthy, and judgments were typically complex.[2] Most approvals required that extensive trackage rights be granted others to ensure competition, and/or that smaller satellite roads be included to guarantee service.

In contrast, proposed mergers in the post-Staggers' era have principally been end-to-end unions,[3] in which the proponents have cited significant economies through the elimination of freight transfer, the implementation of joint marketing agreements, the consolidation of tracking and maintenance systems, increased runthrough operations and more efficient car utilization. While such economies are certainly available in principle, it is also possible that they may be offset and outweighed by organizational issues related to consoliation and merger. Thus, whether these potential economies produce any net efficiency gains in costs or productivity is unclear.

In this context, the potential efficiency gains from mergers fall into three categories: reduced short-run variable costs; reduced long-run costs due to adjustments in way and structure capital stocks toward allocatively efficient levels; and enhanced service quality. Cost functions provide a useful framework in which these efficiencies can be studied.

Consider a total cost (TC) function for railroad h in year t, in which short-run operating and long-run total costs are related as follows:

$$TC_{ht} = SRVC[y_{ht}, w_{ht}, \tau_{ht}, t; K_{ht}] + P_{K,ht}K_{ht} \quad (1)$$

where SRVC is a short-run variable cost function with capital K_{ht} fixed in the short-run, y_{ht} is ton-miles of revenue freight,[4] w_{ht} is a vector of prices for variable inputs (labor, equipment, fuel, materials & supplies),[5] τ_{ht} is a vector of technological factors (route miles, average length of haul, traffic mix), t is a vector of time-related variables (an annual time counter to accommodate disembodied technical change, as well as time counters for years since deregulation occurred, and years since firm h was last involved in a major merger), and $P_{K,ht}$ is an ex ante cost of capital for firm h.

Note that the cost specification in (1) does not include measures of a given railroad's network configuration -- factors that embrace the logistics of a firm's route structure such as end-to-end vs. hub-and-spoke, its track grade, population density over routes, etc. It is of course true that such factors may be important in specifying cross-sectional variations in variable costs. However, specifying a cost function that fully captures these effects would involve introducing a very large number of right-hand side variables into the analysis, many of which would suffer from significant measurement errors. As a practical matter, therefore, we assume that these poorly measured and/or unobserved network configuration effects are invariant for a given railroad over time, and we will attempt to capture the aggregate impact of such network effects on costs by introducing a series of firm-specific dummy variables into the cost function and input share equations; these are discussed further in Section IV below.

Other important measurement issues concern modeling the effects of changes in service quality, and the effects of adjustment toward optimal capital stocks. Service quality measures are typically market specific; for example, Levin and Weinberg [1979] employed market shares in point-to-point markets, while Harris and Winston [1983] used the mean and variance of transit time for specific point-to-point shipments. In both these studies, the network size and capital stocks were implicitly held fixed. This study relies upon firm-level data, rather than point-to-point shipment data. Although we will model the effects of capital adjustment toward optimal long-run levels, we will not be able to account for service quality improvements. In a related paper we quantify the cost savings and capital adjustments of the past fifteen years.[6] In this paper we focus on the issue of measuring the differential effects of mergers and deregulation upon costs and productivity in the rail industry.

We analyze this issue by utilizing a short-run translog variable cost function, written as

$$\ln C_v = A_0 + \sum_{i=1}^{n} A_i \ln w_i + B_1 \ln y + \sum_{j=1}^{m} C_j \ln t_j + D_1 t + M_1 t^m + R_1 t^r$$

$$+ .5 \sum_{i=1}^{n} \sum_{c=1}^{n} AA_{ic} \ln w_i \ln w_c + .5\ BB_{11} (\ln y)^2 + \sum_{i=1}^{n} AB_{i1} \ln w_i \ln y$$

$$+ \sum_{i=1}^{n} \sum_{j=1}^{m} AC_{ij} \ln w_i \ln t_j + \sum_{i=1}^{n} BC_{j1} \ln t_j \ln y + .5 \sum_{j=1}^{m} \sum_{k=1}^{m} CC_{jk} \ln t_j \ln t_k$$

$$+ \sum_{i=1}^{n} AD_{i1} t \cdot \ln w_i + \sum_{i=1}^{n} AD_{i2} t^{r} \cdot \ln w_i + \sum_{i=1}^{n} AD_{i3} t^{m} \cdot \ln w_i + AD_4 MDUM$$

$$+ .5\, D_{11} t^2 + .5\, M_{11} t^{m2} + .5\, R_{11} t^{r2} \qquad (2)$$

where C_v is variable costs, $C_v = \sum_i w_i x_i$, w_i and x_i are input prices and quantities, respectively, $i = 1,\ldots,n$, y is output, the t_j are technological variables, $j = 1,\ldots,m$, t is an annual time counter having the value one in the first year of the sample (1974), t^m is an annual time counter on the number of years since the company has last been involved in a merger, MDUM is a dummy variable taking on a value of one if the railroad was involved in a merger in that year (else it is zero), and t^r is another annual time counter measuring years since the 1978 deregulation occured (1979=1,...,1986=8). According to this specification, mergers may have a one time positive "adjustment cost" effect on costs as measured by AD_4 (the coefficient on the MDUM dummy variable), but after that costs can diminish over time, depending on the values of the M_1 and M_{11} parameters.

For the translog restricted cost function to be well-behaved, it is necessary that symmetry and homogeneity restrictions be imposed. These take the form

$$AA_{ic} = AA_{ci},\ CC_{jk} = CC_{kj}\ ,\quad i,c = 1,\ldots,n;\ j,k = 1,\ldots,m$$

$$\sum_{i=1}^{n} A_i = 1,\quad \sum_{i=1}^{n} AA_{ic} = \sum_{i=1}^{n} AB_{i1} = \sum_{i=1}^{n} AC_{ij} = \sum_{i=1}^{n} AD_{ip} = 0,$$

$$i,c = 1,\ldots,n;\ j = 1,\ldots,m;\ p = 1,2,3. \qquad (3)$$

Logarithmically differentiating (2) with respect to input prices and using Shephard's Lemma yields n cost share equations, only n-1 of which are linearly independent. The i^{th} share equation is of the form

$$\frac{\partial \ln C_v}{\partial \ln w_i} = \frac{w_i x_i}{C_v} = A_i + \sum_{c=1}^{n} AA_{ic} \ln w_c + \sum_{j=1}^{n} AC_{ij} \ln t_j + AB_{i1} \ln y$$

$$+ D_{i1} t + D_{i2} t^r + D_{i3} t^m , \; i = 1,\ldots,n. \qquad (4)$$

We analyze productivity growth by differentiating the cost function (2) with respect to time, which yields the following expression for the rate of cost diminution (the negative of the rate of productivity growth):

$$\frac{\partial \ln C_v}{\partial t} = D_1 + M_1 \cdot FM + R_1 \cdot FD + \sum_{i=1}^{n} (AD_{i1} + AD_{i2} \cdot FD + AD_{i3} \cdot FM) \cdot \ln w_i$$

$$+ D_{11} t + M_{11} \cdot FM \cdot t^m + R_{11} \cdot FD \cdot t^r, \qquad (5)$$

where FM and FD are dummy variables indicating whether the firm was ever involved in a merger ($FM=1$ if yes), and if the observation is in a year after deregulation has occurred ($FD=1$ if yes). According to (5), productivity growth differs depending on whether the observation comes from a regulated time period (in which case the R_1, AD_{i2} and R_{11} terms are zero) or a deregulated era (these terms become non-zero), and, for firms having been involved in a merger or acquisition, the number of years since the last merger (the M_1, AD_{13} and M_{11} terms become non-zero).

The effects of deregulation and mergers on costs and productivity can be partially quantified as follows. If deregulation and mergers have no effects on costs and productivity, then $M_1 = R_1 = AD_{i2} = AD_{i3} = AD_4 = M_{11} = R_{11} = 0$, $i = 1,\ldots,n$. But if deregulation has differential effects on costs and productivity while mergers do not, then $M_1 = AD_{i3} = AD_4 = M_{11} = 0$ and some of the R_1, AD_{i2}, and $R_{11} \neq 0$, $i = 1,\ldots,n$. Finally, if mergers have differential effects on costs and productivity while deregulation does not, then $R_1 = AD_{i2} = R_{11} = 0$, and some of the M_1, AD_{i3}, AD_4 and $M_{11} \neq 0$. Hence, within this framework, one can distinguish in part the effects on costs and productivity of deregulation from those directly due to mergers and acquisitions.

III. DATA SOURCES AND ECONOMETRIC ISSUES

The estimation of a short-run variable cost function and its input demand equations requires the following firm-specific data: short-run variable costs by input, input factor prices, output measures and output composition; technological factors; and way & structures capital stock. We now provide a brief overview of our data set and its construction; further details are given in Vellturo [1989].

We begin with a discussion of firms in our sample. Fifty-six railroads had Class I status in 1974, but only twenty-one reported data in 1986.[7] From these systems, 27 railroads were deemed to have sufficiently accurate data to be included in the sample. A large number of railroads lost Class I status, others had incomplete bond histories (which made it impossible to generate correct capital equipment costs), while others went bankrupt for significant portions of the sample and their data was deemed unreliable. Since certain roads ceased to exist upon consolidation into other new systems, the panel is not balanced; if it were balanced, our panel would have 351 observations (27 firms times 13 years), but for us only 229 observations are available. The names of the firms used in our sample, and their abbreviations, are given in Table 1 below.

Table 1

US CLASS I RAILROADS, 1974-86

Railroad/System	Abbreviation	Years Observed In Data Set
Ann Arbor	AA	
Atchison, Topeka & Santa Fe	ATSF	74-86
Baltimore & Ohio	BO	
Bessemer, Lake Erie	BLE	
Boston & Maine	BM	
Burlington Northern	BN	74-79
Central of New Jersey	CNJ	
Chesapeake & Ohio	CO	
Chicago, Northwest Transit	CNWT	74-86
Chicago, Milwaukee & St. Paul	MILW	
Colorado Southern	CS	74-81
Delaware & Hudson	DH	
Denver, Rio Grande Western	DRGW	74-85
Detroit, Toledo Shore Line	DTS	
Detroit, Toledo & Ironton	DTI	
Duluth, Missabe & Iron Range	DMIR	
Elgin, Joliet & Eastern	EJE	
Erie Lackawanna	EL	
Florida East Coast	FEC	
Fort Worth, Denver	FWD	74-81
Grand Trunk & Western	GTW	75-86
Illinois Central Gulf	ICG	74-86
Kansas City Southern	KCS	74-86
Lehigh Valley	LV	
Louisville & Nashville	LN	
Missouri-Kansas-Texas	MKT	74-86
Missouri Pacific	MP	74-82
Norfolk & Western	NW	74-81
Penn Central	PC	
Pittsburgh, Lake Erie	PLE	
Reading Railroad	RDG	
St. Louis, San Francisco	SLSF	74-79
St. Louis, Southwestern	SSW	
Seaboard Coast Line	SCL	
Soo Line	SOO	74-86
Southern Pacific	SP	74-86
Southern Railway System	SOU	74-81
Union Pacific Railway	UP	74-82
Western Maryland	WM	
Western Pacific	WP	74-82
Consolidated Rail Corp.	CRC	77-86
Clinchfield & Ohio	CLINCH	
Chessie System	CHES1	74-80
Seaboard System	SBD	74-80
CSX Corporation (1981-82)	CSX1	81-82
CSX Corporation (1983-86)	CSX2	83-86
Burlington Northern - St. Louis System	BNSL	80-81
Burlington Northern System	BNSYS	82-86
Union Pacific System	UPSYS	83-86
Norfolk-Southern Corporation	NS	82-86

In order to adjust for the effects of inflation, we estimate the cost function in which all dollar magnitudes are expressed in real dollars.[8] Thus variable costs, input expenses and prices, and the capital stock are deflated and measured in constant 1971 dollars.

For each firm in the sample, data series on operating expenses as reported by Class I railroads annually to the ICC required two significant adjustments before they could be used as variable costs. First, equipment costs are reported to the ICC through depreciation of the rolling stock. Instead, we constructed a rental price for equipment, allowing both for economic depreciation and the cost of floating equipment trusts, the most common form of rolling stock financing. We then multiplied this rental price measure times an estimate of the replacement value of the rolling stock for each firm, to obtain an estimate of the annual cost of equipment services. Second, prior to 1983, rail firms expensed maintenance on way & structures capital in the year such maintenance was undertaken.[9] Instead, we removed these maintenance charges from variable costs and added them to net capital investment in the construction of the replacement value of each road's way & structure stock.

For factor prices, our goal was to obtain as disaggregated a series as possible, cross-sectionally. For fuel and "materials and supplies", we took published regional prices by the ICC and assigned them (or a weighted sum of them) to individual railroads; thus for differing railroads operating entirely in the same region, the price of materials is the same for the various firms, but for firms operating in more than one region, the materials price is firm-specific.[10] For wage rates, a Divisia index was constructed for each road, based upon labor compensation data for 78 wage categories.[11] The equipment factor price represents the user cost of capital, and incorporates firm-specific costs of capital, as we noted earlier.

A single output measure, total ton-miles of revenue and non-revenue freight, was deemed the most reliable measure available. Since AMTRAK handles the overwhelming majority of non-commuter passenger traffic, it is reasonable to use only freight as a measure of output.[12] Another multiproduct option we considered involved freight output by type of traffic (e.g., coal, agriculture and manufactured goods). Unfortunately, railroads only account for tons of traffic by commodity type, not ton-miles. As a result, we included percent of tons carried for each commodity type (coal, agriculture and other) as technological factors in the short-run cost function specification.[13] Other technical factors include average length of haul (total ton miles divided by tons carried), and miles of main and secondary track owned and operated by each system.

As we noted earlier, way & structures capital stocks for each road have undergone significant accounting restructuring at the ICC over the sample period. Independent of these changes is the fact that ICC measures of capital are book values. Constant-dollar capital stocks, measured using reproduction value concepts, is clearly a more appropriate capital measure in assessing the effect a marginal unit of capital has on variable costs. Fortunately, Nelson [1974] has provided estimates of reproduction values of way & structure stocks for each major Class I road in 1971. Using Nelson's figures and subsequent investment data, a geometric depreciation schedule based upon capital lifetimes, and maintenance expenditures, we have constructed estimated capital stocks in constant dollars for each major Class I road. Further details of this procedure are given in Vellturo [1989]. Sample means and standard deviations of variables in our data set are given in Table 2 below.

With this information as background, we now move on to a discussion of some econometric issues associated with the estimation of this cost function.

Table 2

Means and Standard Deviations of Variables Used
in Econometric Analysis of Railroad Costs, 1974-1986

Variable	Units	Mean	Std. Dev.	Min Value	Max Value
Variable Costs	Billion$*	1.141	1.148	0.019	5.040
Price of Labor	$/Hour*	9.677	2.843	5.390	17.740
Price of Equip	Index**	0.396	0.131	0.190	0.674
Price of Fuel	Index***	1.637	0.742	0.684	2.844
Price of M+S	Index***	1.173	0.260	0.652	1.495
Ton Miles	Billion	45.245	41.876	1.910	203.000
Percent Ag	% Points	19.117	9.025	6.298	69.175
Percent Coal	% Points	26.403	17.816	0.291	79.378
W&S Capital	Billion$	1.906	1.923	0.118	8.303
ALOH	1000 Miles	0.392	0.145	0.173	0.780
Route Mileage	1000 Miles	7.930	6.428	0.543	25.810
Time (Year)	Year - 1973	6.476	3.594	1.000	13.000
Labor Expenses	Billion$*	0.460	0.464	0.004	1.940
Equip Expenses	Billion$*	0.373	0.416	0.003	2.175
Fuel Expenses	Billion$*	0.116	0.123	0.003	0.547
M+S Expenses	Billion$*	0.192	0.205	0.001	1.137
Labor Share		0.397	0.081	0.105	0.664
Equip Share		0.333	0.091	0.068	0.585
Fuel Share		0.105	0.039	0.036	0.301
M+S Share		0.165	0.064	0.008	0.459

Notes:

*Variable costs, the price of labor, and input expenses are in constant 1971$
**Indexed to 1977 = 1.000
***Indexed to 1971 = 0.100

IV. SPECIFICATION OF COST FUNCTION AND ISSUES IN ECONOMETRIC IMPLEMENTATION

We begin with a discussion of the stochastic specification of our equation system, one that explicitly incorporates the notion that variables not observed by the econometrician are important to the economic agent; in turn, the actions of this economic agent on input demands and variable costs reflect cost minimization in a coherent manner. Specifically, we assume that the cost equation and its associated input share equations have additive error component structures of the following form

$$\ln C_{v,ht} = F(w, y, t, \tau, K; \beta)|_{ht} + \epsilon_{ht} \quad (6)$$

$$S_{iht} = G(w, y, t, \tau, K; \beta)|_{ht} + \mu_{iht}, \; i = 1,\ldots,n, \quad (7)$$

where the variables have their previous definitions and β represents the vector of parameters associated with the estimated equations. We decompose each of these error terms into three components: a firm-specific error (α_h and α_{ih}); an error that exhibits first order autocorrelation within a given equation (δ_t and γ_{it}), but no error autocorrelation across equations; and a normally distributed disturbance term that is only contemporaneously correlated across equations (ϕ_{ht} and ω_{iht}). These errors are therefore written as

$$\epsilon_{ht} = \alpha_h + \delta_t + \phi_{ht}, \qquad h = 1,\ldots,H, \; t = 1,\ldots,T \quad (8)$$

$$\mu_{iht} = \alpha_{ih} + \gamma_{it} + \omega_{iht}, \qquad i = 1,\ldots,n. \quad (9)$$

To motivate this stochastic specification, we begin by considering the origin of the firm-specific error terms (α_h and α_{ih}). We interpret these firm-specific disturbances as reflecting unobserved fundamental network differences among Class I railroads (e.g., the spacial configurations of their routes, whether networks are primarily hub-and-spoke, end-to-end, etc.). Since it is reasonable to assume that this network configuration effect is fixed over time for a given railroad, we can eliminate this firm-specific error component by introducing indicator variables for each firm. We assume that these network attributes are known to the railroads but not observed by the econometrician, that the underlying network configuration influences input utilization decisions by firms, and therefore that these unobserved effects enter into their cost-minimizing behavior.

We implement these assumptions by introducing dummy variables into the linear term of the input share equations, and, since the share equations are derived as logarithmic derivatives from the variable cost function, we also introduce interactive dummy*ln w_i slope coefficients into the variable cost function equation, constraining their coefficient values to equal those in the input share equations; our interpretation of the unobserved variables therefore implies that cross-equation restrictions should be imposed on the fixed-effect parameters.[14]

Thus the coefficients given in the cost (2) and input share (4) equations should be interpreted as follows:

$$A_0 = A_0' + F_h, \quad h = 1,\ldots,H-1 \tag{10}$$

$$A_i = A_i' + F_{ih}, \quad i = 1,\ldots,n; \quad h = 1,\ldots,H-1, \tag{11}$$

where A_0' and A_i' respectively represent the intercept and linear coefficients on the input price variable for the base railroad (denoted H); F_h is a zero-one intercept dummy for railroad h in the cost function, and F_{ih} is both a zero-one intercept dummy for railroad h in the i^{th} share equation *and* a multiplicative dummy variable on the ln w_i term in the variable cost function equation.

In the context of mergers, we define new firm-specific fixed effect dummy variables for merged firms; obviously, the component firms of the merger are dropped from the sample once a merger occurs. Given the nonlinear nature of the translog variable cost specification (2), it is not possible to relate analytically the fixed effects of merged firms to those of their pre-merger constituent firms. As a result, we can only perform a partial parametric test on the effects of mergers. In particular, although appropriate restrictions can be tested on merger dummy variables and merger-related slope coefficients, it is not possible to establish and therefore test for relationships between the fixed effect of a merged firm and those of its pre-merger constituent firms.

Given this fixed effect specification, appropriate adding-up conditions must also be imposed on the fixed effects coefficients. These turn out to take the form

$$F_{nh} = -\sum_{j=1}^{n-1} F_{jh}\ , \quad h = 1,\ldots,H-1, \text{ and } \quad F_{n0} = 1 - \sum_{j=1}^{n-1} F_{j0} \qquad (12)$$

Returning to our stochastic specification, we introduce intra-equation intertemporal effects by permitting the δ_t and γ_{it} terms to follow first-order autoregressive processes. Although equal across firms, we specify that the first order autoregressive parameter in the cost function disturbance term δ_t differs from that in the share equation disturbance terms γ_{it}, but to ensure adding-up consistency, we also specify that the autoregressive parameter for each share equation is equal across shares.[15]

Cross-equation contemporaneous correlation of the ϕ_{ht} and ω_{iht} terms is expected, due to the adding up of the share equations. Therefore, we specify that the n-element disturbance vector consisting of the ϕ_{ht} and n-1 ω_{iht} terms is independent and multivariate normally distributed, with mean vector zero and covariance matrix Ω_{ht}. Finally, on the basis of preliminary examination of residuals, it was determined that in both the share and cost function equations, heteroskedasticity occurred, with the variance of residuals being positively related to the $(\ln y)_{ht}$. To transform the model so that the disturbance terms became homoskedastic, we therefore divided all variables by the square root of $(\ln y)_{ht}$.

One other matter deserves particular attention in the context of stochastic specification, and that concerns the exogeneity of output (ln y), ALOH and the output composition variables %AG and %COAL. Because of the rate-setting freedom introduced by the Staggers Act, it is reasonable to ask whether output and its composition should be treated as being exogenous, particularly after 1978. If they are endogenous, output and its components may well be correlated with the cross-equation error terms, so that estimates based on the assumption of exogeneity of these regressors could be biased and inconsistent. To test for the validity of the exogeneity assumption, we have utilized the specification test procedure due to Hausman [1978] and Hausman et al. [1988].

Specifically, insofar as output and its components are determined endogenously through the profit-maximizing behavior of railroads, they should be related to demand variables that do not enter the cost function. Consequently, in implementing the Hausman specification test, we utilized as instruments appropriate firm-specific demand-related variables, including coal production, mine-mouth prices, oil rates, farm income and value of shipments from manufacturing.[16] We now turn to a discussion of empirical findings.

V. ECONOMETRIC IMPLEMENTATION AND RESULTS FROM ESTIMATION

We begin by implementing a system version of the Hausman specification test, estimating the equation system consisting of the variable cost function (2) and n-1 of the cost share equations (4), first by the 3SLS procedure assuming that ln y, ALOH, %AG, %COAL and their transforms are endogenous, and then by maximum likelihood (ML) under the assumption that all regressors are uncorrelated with error terms. Note that under the null hypothesis, ML estimation is efficient while 3SLS is consistent, and if the alternative hypothesis is true, then only the 3SLS estimates are consistent.[17] The χ^2 test statistic corresponding to the null hypothesis that $\beta_{3SLS} = \beta_{ML}$ is 1252.988, which is much larger than the critical value with 38 degrees of freedom at any reasonable significance level. We conclude that ln y, ALOH,

%AG and %COAL are not exogenous variables, but instead are endogenous. All subsequent estimation results that we report are therefore based on the assumption that these output-related variables are endogenous (i.e., we only report the 3SLS estimates).

Using the residuals from the 3SLS model, we estimated a common autoregressive parameter for the three share equations, and another autoregressive parameter for the cost function; the null hypothesis that these two autoregressive parameters were simultaneously equal to zero was not rejected at usual significance levels.[18] Hereafter we set these autocorrelation coefficients to zero.

The 3SLS estimated model had one additional drawback, in that curvature restrictions involving way & structures capital were frequently violated.[19] Inspection of our parameter estimates revealed that the CC_{11} coefficient was positive but very imprecisely estimated. We therefore constrained this parameter to zero.[20] Once this restriction was imposed, curvature restrictions were satisfied at 195 of our 229 observations. Parameter estimates and t-statistics (based on heteroskedas-ticity-robust standard errors) for the common parameters are given in Table 3, firm-specific effects estimates are presented in Table 4,[21] and the merger-deregulation parameters and test statistics are highlighted in Table 5.[22]

We begin our discussion of results by examining the differential effects on costs and productivity growth of deregulation from those of mergers and acquisitions. In Table 5 we summarize time-related parameter estimates for a model in which both deregulation and merger variables are incorporated (Model I), a model in which deregulation variables are excluded but merger variables are included (Model II), and a model in which the merger variables are

Table 3

3SLS PARAMETER ESTIMATES FOR RESTRICTED TRANSLOG SHORT-RUN COST FUNCTION
("t-stat" is the Ratio of Parameter Estimate to its Asymptotic Standard Error)
US Class I Railroads, 1974-1986

Parameter	Variable	Estimate	t-Stat	Parameter	Variable	Estimate	t-Stat
B_1	y	5.5726	3.09	AA_{11}	$P_L \cdot P_L$	0.1654	2.89
C_1	K	1.3454	1.26	AA_{22}	$P_E \cdot P_E$	0.1530	3.26
C_2	ALOH	-3.1596	-1.20	AA_{33}	$P_F \cdot P_F$	0.0791	8.04
C_3	MILES	-7.1830	-3.00	AA_{44}	$P_{M+S} \cdot P_{M+S}$	0.1253	2.10
C_4	% AG	4.6114	2.88	AA_{12}	$P_L \cdot P_E$	-0.1298	-4.52
C_5	% Coal	-2.0086	-3.39	AA_{13}	$P_L \cdot P_F$	-0.0128	-0.80
AB_{11}	$y \cdot P_L$	0.0442	1.68	AA_{14}	$P_L \cdot P_{M+S}$	-0.0228	-0.40
AB_{21}	$y \cdot P_E$	-0.2042	-10.00	AA_{23}	$P_E \cdot P_F$	0.0064	0.35
AB_{31}	$y \cdot P_F$	0.0766	3.69	AA_{24}	$P_E \cdot P_{M+S}$	-0.0296	-0.80
AB_{41}	$y \cdot P_{M+S}$	0.0835	2.12	AA_{34}	$P_F \cdot P_{M+S}$	-0.0728	-3.39
AC_{11}	$P_L \cdot K$	0.0860	1.26	BC_{11}	$K \cdot y$	0.6115	3.01
AC_{21}	$P_E \cdot K$	0.0567	0.90	BC_{12}	ALOH·y	0.0375	0.11
AC_{31}	$P_F \cdot K$	0.0519	1.89	BC_{13}	MILES·y	-0.5425	-1.95
AC_{41}	$P_{M+S} \cdot K$	-0.1945	-3.20	BC_{14}	%AG·y	0.1882	0.98
AC_{12}	$P_L \cdot$ALOH	-0.0032	0.07	BC_{15}	%COAL·y	0.0130	0.21
AC_{22}	$P_E \cdot$ALOH	0.1228	3.11	BB_{11}	$y \cdot y$	0.0899	0.25
AC_{32}	$P_F \cdot$ALOH	0.0219	0.61	CC_{11}	$K \cdot K$	0.0000	
AC_{42}	$P_{M+S} \cdot$ALOH	-0.1415	-1.90	CC_{12}	ALOH·K	-0.5618	-2.25
AC_{13}	$P_L \cdot$MILES	0.0437	1.24	CC_{13}	MILES·K	-0.4126	-2.27
AC_{23}	$P_E \cdot$MILES	0.0095	0.31	CC_{14}	%AG·K	0.5892	3.47
AC_{33}	$P_F \cdot$MILES	-0.0226	-1.53	CC_{15}	%COAL·K	-0.1630	-3.85
AC_{43}	$P_{M+S} \cdot$MILES	-0.0306	-0.72	CC_{22}	ALOH·ALOH	1.0303	2.27
AC_{14}	$P_L \cdot$%AG	0.0487	1.87	CC_{23}	ALOH·MILES	0.2689	0.65
AC_{24}	$P_E \cdot$%AG	-0.0404	-2.02	CC_{24}	ALOH·%AG	-0.5902	-2.92
AC_{34}	$P_F \cdot$%AG	0.0023	0.85	CC_{25}	ALOH·%COAL	-0.1494	-1.37
AC_{44}	$P_{M+S} \cdot$%AG	-0.0106	-0.23	CC_{33}	MILES·MILES	1.0130	2.86
AC_{15}	$P_L \cdot$%COAL	-0.0144	-2.11	CC_{34}	MILES·%AG	-0.6779	-3.13
AC_{25}	$P_E \cdot$%COAL	0.0161	3.16	CC_{35}	MILES·%COAL	0.1947	2.52
AC_{35}	$P_F \cdot$%COAL	-0.0205	-3.63	CC_{44}	%AG·%AG	0.1160	0.58
AC_{45}	$P_{M+S} \cdot$%COAL	0.0188	2.00	CC_{45}	%AG·%COAL	-0.0759	-1.17
AD_{T1}	$P_L \cdot$TIME	0.0009	0.27	CC_{55}	%COAL·%COAL	-0.0199	-1.30
AD_{T2}	$P_E \cdot$TIME	0.0009	0.27	DT	TIME	-0.1284	-4.10
AD_{T3}	$P_F \cdot$TIME	-0.0125	-5.13	DTT	TIME·TIME	0.0509	5.37
AD_{T4}	$P_{M+S} \cdot$TIME	0.0107	2.50	M_1	t^m	-0.0718	-2.40
R_1	t^r	-0.0909	-2.96	M_{11}	t^{m2}	0.0097	2.29
R_{11}	t^{r2}	-0.0537	-6.25	AD_4	MDUM	-0.0724	-3.14
AD_{12}	$P_L \cdot t^r$	-0.0064	-1.35	AD_{13}	$P_L \cdot t^m$	-0.0029	-0.83
AD_{22}	$P_E \cdot t^r$	0.0009	0.20	AD_{23}	$P_E \cdot t^m$	0.0009	0.20
AD_{32}	$P_F \cdot t^r$	0.0108	3.52	AD_{33}	$P_F \cdot t^m$	0.0108	3.52
AD_{42}	$P_{M+S} \cdot t^r$	-0.0053	-0.94	AD_{43}	$P_{M+S} \cdot t^m$	-0.0053	0.94

NOTES: Prices, output quantity, K, ALOH, MILES, %AG, %COAL are all logarithmically transformed. TIME, t^m (years since last merger) and t^r (years since deregulation) are all in natural units. Standard error estimates employ the Halbert White [1980] heteroskedasticity-robust computation.

Table 4

3SLS PARAMETER ESTIMATES FOR FIRM-SPECIFIC COST AND FACTOR SHARE TERMS
TRANSLOG SHORT-RUN COST FUNCTION, US CLASS I RAILROADS, 1974-1986
(Asymptotic t-statistic based on robust standard errors)

	Cost Function		Labor Term		Equipment Term		Fuel Term	
Firm	Estimate	t-Stat	Estimate	t-Stat	Estimate	t-Stat	Estimate	t-Stat
Base Firm:								
ATSF	21.027	2.43	-0.708	1.92	1.500	5.57	0.003	0.02
Other Firms:								
BN	-0.095	-0.47	-0.033	-0.73	-0.027	-0.68	-0.038	-2.04
CNWT	-1.724	-6.17	0.238	3.70	-0.130	-2.63	0.128	3.05
CS	-2.906	-3.62	0.266	1.19	-0.268	-1.33	0.404	3.59
DRGW	-1.964	-3.50	0.357	2.25	-0.220	-1.47	0.256	2.96
FWD	-4.152	-5.41	0.594	2.85	-0.485	-2.52	0.413	4.50
GTW	-2.062	-3.26	0.498	2.73	-0.241	-1.49	0.249	2.41
ICG	-0.828	-3.69	0.107	1.88	-0.023	-0.74	0.081	1.86
KCS	-2.213	-3.56	0.355	2.00	-0.114	-0.71	0.229	2.56
MKT	-1.994	-3.45	0.312	1.92	-0.093	-0.64	0.237	3.07
MP	-0.734	-4.46	0.090	2.30	-0.008	-0.25	0.054	1.91
NW	-0.020	-0.09	0.073	1.52	0.081	2.09	0.026	0.64
SLSF	-1.473	-4.18	0.236	2.53	-0.158	-1.90	0.163	3.35
SOO	-1.933	-4.73	0.259	2.29	-0.165	-1.61	0.150	2.65
SPTC	0.152	1.52	-0.044	-2.10	0.072	4.01	-0.068	-3.79
SOU	-0.022	-0.10	0.016	0.32	0.097	2.70	0.038	0.85
UP	0.054	0.55	0.005	0.24	0.007	0.40	0.011	0.96
WP	-2.459	-3.81	0.323	2.11	-0.350	-2.42	0.204	3.07
CRC	0.025	0.06	0.068	0.66	-0.190	-2.11	-0.054	-0.99
CHES1	0.116	0.34	0.073	1.10	0.069	1.52	0.003	0.05
SBD	-0.347	-1.29	0.009	0.17	0.138	4.06	0.009	0.16
BNSL	-0.147	-0.57	-0.039	-0.74	-0.011	-0.22	-0.024	-1.09
BNSYS	-0.014	-0.05	-0.070	-1.22	0.018	0.32	-0.052	-2.20
CSX1	0.194	0.44	-0.068	-0.74	0.202	2.65	-0.073	-1.18
CSX2	0.292	0.71	-0.052	-0.59	0.147	2.00	-0.077	-1.32
NS	0.502	1.74	-0.055	-0.89	0.203	3.95	-0.049	-1.08
UPSYS	0.305	1.46	-0.067	-1.40	0.131	2.94	-0.057	-2.34

Notes: BNSL and BNSYS represent the Burlington Northern System over two phases, 1979-80 (before the acquisition of FWD and CX), and 1981-86 (after the acquisition). CSX1 and CSX2 represent the two phases of the CSX merger, 1981-82 and 1983-86.

Table 5

3SLS ESTIMATES OF DEREGULATION AND MERGER COEFFICIENTS
ALTERNATIVE MODEL SPECIFICATIONS, CLASS I RAILROADS, 1974-86
(Absolute Value of Coefficient Estimate to
Heteroskedasticity-Robust Asymptotic Standard Error in Parentheses)

Parameter	Variable	Model I	Model II	Model III
M_1	t^m	-.07183 (2.4008)	-.06848 (2.2334)	.00000
M_{11}	t^{m2}	.00973 (2.2950)	.00959 (2.1659)	.00000
AD_4	MDUM	-.07238 (3.1368)	-.06666 (2.8824)	.00000
AD_{13}	$t_m \cdot w_1$	-.00291 (0.8319)	-.00335 (??????)	.00000
AD_{23}	$t_m \cdot w_2$	.00055 (0.1937)	.00144 (0.4964)	.00000
AD_{33}	$t_m \cdot w_3$	.00150 (0.9775)	.00440 (2.2175)	.00000
AD_{43}	$t_m \cdot w_4$	.00087 (0.1504)	-.00249 (0.3587)	.00000
	* *			
R_1	t^r	-.09087 (2.9597)	.00000	-.06777 (2.2322)
R_{11}	t^{r2}	-.05373 (6.2501)	.00000	-.05110 (5.7998)
AD_{12}	$t_r \cdot w_1$	-.00640 (1.3516)	.00000	-.00716 (??????)
AD_{22}	$t_r \cdot w_2$	.00089 (0.1998)	.00000	.00096 (0.2264)
AD_{32}	$t_r \cdot w_3$	.01077 (3.5215)	.00000	.01093 (3.6587)
AD_{42}	$t_r \cdot w_4$	-.00526 (0.9360)	.00000	-.00473 (0.8669)
E'HH'E		481.779	581.522	493.849

excluded but the deregulation variables are included (Model III). As is seen in the Model I column of Table 5, the linear and the squared time counter estimates for the merger (M_1 and M_{11}, see top panel) parameters are significant, as are both deregulation (R_1 and R_{11}, see bottom panel) parameters. Both M_1 and R_1 are negative, but R_1 is about 27% larger in absolute value; interestingly, M_{11} is positive and small (.0097), while R_{11} is negative and quite large (-0.0537). Together with the substantial negative estimate of the merger dummy variable MDUM (equals one if the firm was involved in a merger that year), these results imply that mergers appear to have had a cost reducing impact, but that in our sample this effect diminishes with time; by contrast, while the negative effect of deregulation on costs is also substantial, the large negative estimate of R_{11} indicates that in our sample, the cost-reducing effects of deregulation increased over time. In brief, the cost-reducing effects of mergers are smaller and shorter-lived than are those due to deregulation.

The cross-product terms involving the time and factor price variables are also of particular interest here. From Table 3 we see that the estimate of AD_{T1} is positive, very small, and statistically insignificant, implying that in the absence of deregulation and mergers, the effects of time on labor demand are approximately neutral. Recall that estimates of AD_{13} and AD_{12} indicate whether the effects of time on the labor share are different in the presence of deregulation (AD_{12}) or mergers (AD_{13}). As is seen in Table 5, deregulation has a non-neutral labor-saving effect on costs, but so too do mergers; the estimated effect of deregulation (-0.00640) is more than twice as large as that due to mergers (-0.00291). The precision with which these labor bias effects are measured is not great, although the 1.35 t-value on the deregulation labor coefficient is larger than the 0.83 t-value on the merger labor coefficient.

We then tested Model II (deregulation doesn't matter) as a special case of Model I; the quasi-likelihood ratio test statistic is 99.74, which is greater than the chi-square test statistic with five degrees of freedom at any reasonable significance level. We also tested Model III (mergers don't matter) as a special case of Model I. This too was rejected, although not nearly as decisively; the chi-square test statistic with six degrees of freedom is 12.07, which implies rejection of the null hypothesis at all significance levels less than 0.06043.[23] We conclude that both mergers and deregulation had significant cost-reducing effects on railroads in the US from 1974 to 1986, although our estimates of effects of deregulation are measured more decisively.

To obtain some idea of the relative quantitative magnitudes of mergers and deregulation on costs and productivity, we simulated costs of all firms involved in mergers since 1980,[24] we predicted what their costs would have been had they not merged,[25] and we predicted what costs would have been had deregulation not occurred. By 1986, of the total cumulative cost savings achieved by merged firms since 1980, that due to deregulation is estimated at 91%. while those due directly to mergers is estimated to be 9%. We conclude, therefore, that while mergers and deregulation both had substantial cost-reducing effects in our 1974-86 sample, those due to deregulation are much larger.

VI. CONCLUDING REMARKS

In this paper we have sought to disentangle the effects of deregulation on rail costs from those directly attributable to mergers and acquisitions. Our empirical analysis has employed a translog short-run variable cost function, based on an unbalanced panel data set of annual observations for major US Class I railroads from 1974 to 1986. In our econometric implementation, we specify that the fixed effects reflect variables not observed by the econometrician, but are known and important to the cost-minimizing firm; this interpretation implies that the traditional fixed effect model should be modified to incorporate cross-equation restrictions on the fixed effect parameters.

Our first empirical finding is that both deregulation and mergers contributed significantly to cost savings; hypotheses of the form that "deregulation has no effect" and "mergers have no cost effects" were both strongly rejected, although the former was rejected much more decisively than the latter.

Second, although mergers accounted for some of the cost savings observed in firms that have consolidated during our sample period, by 1986 the bulk of the accumulated cost savings achieved by merged firms since 1980 appears to be due to deregulation (about 91%). Moreover, the cost-reducing effects of mergers are short-lived than are those due to deregulation.

Third, in terms of factor biases, we find that both deregulation and mergers resulted in a substantial labor-saving bias; although our point estimate of this labor-saving effect for deregulation is more than twice as large as that for mergers, we were not able to estimate the magnitude of these bias coefficients precisely.

These findings suggest that while mergers contributed modestly to the cost and productivity improvements that occurred during our 1974-86 sample period, deregulation has also had an enormous impact; indeed, its impact appears to have been much larger.

FOOTNOTES

[1]For example, major mergers such as the formation of the Penn-Central railroad (from the New York Central and Pennsylvania roads), and the Burlington Northern (from the Great Northern, Northern Pacific, and Burlington Railroads) all involved consolidations of parallel trackage.

[2]As an example, the Burlington Northern application was originally presented to the ICC in 1959. Final approval was granted eleven years and three resubmissions later. The infamous Penn-Central application spanned seven years. For details, see Saunders [1978].

[3]One notable exception involved the proposed consolidation of the Southern Pacific and Atchison, Topeka and Santa Fe Railways, brought to the ICC in 1984. Both roads had extensive trackage that ran virtually parallel to one another throughout the southwest. The merger was rejected.

[4]A number of previous studies have specified a multi-product variable cost function with ton-miles of freight and passenger miles treated as distinct outputs (e.g., Friedlaender and Spady [1981]). Since the advent of AMTRAK, however, passenger traffic has comprised less than 1% of total car-miles for Class I railroads. Hence we ignore passenger traffic in this study. An entirely different issue, however, is the distinction with ton-miles by type of traffic, such as coal, agriculture, and manufactured goods. We address this issue later in this paper.

[5]In this study fixed capital refers to "Way and Structures" capital, as it is known in the industry. Such capital includes track, grading, bridges, buildings, etc. Equipment capital (commonly referred to as "rolling stock") is treated as a variable input in the short-run, since an active rental market exists for such equipment.

[6]See Friedlaender, Berndt, Chiang and Vellturo [1990].

[7]During this period, the ICC changed the minimum freight revenue required to be considered a Class I railroad from $1 million to $5 million in 1978, and to $10 million in 1983. Non-Class I railroads are required to file far less detailed cost schedules with the ICC.

[8]To deflate costs, we use the Producer Price Index, taken from the Economic Report of the President.

[9]In 1983, the ICC recognized that this approach led to a severe miscalculation of the replacement value of the way and structure capital for most roads. This form of accounting (known as "betterment accounting") was replaced with a depreciation-based accounting procedure, and way & structure replacement values underwent a one-time adjustment to capture the misclassified maintenance expenditures.

[10]Until 1985, railroads were classified into the east, south or west region, but mergers prompted the ICC to consolidate the east and south regions in 1985.

[11]The relative values of this price index for labor among firms was set equal to relative average values in 1974.

[12]Local transportation authorities handle commuter traffic in nearly every major metropolitan area in the US. The notable exception is the Chicago Northwest Transit Co., which provides a portion of Chicago's rail needs.

[13]An alternative approach involved exploiting the relationship between total ton-miles and tons by commodity type, ton-miles = $\sum_i ALOH_{ri} \cdot Tons_{ri}$ for railroad i for various commodity types i, where ALOH is average length of haul. We estimated this relationship using time series data for each railroad to obtain estimated average lengths of haul by commodity type. Although most estimates were positive, they were very imprecise and we therefore abandoned this effort.

[14]In estimating rail costs, Caves et al. [1985] specified that fixed effects enter the cost function, but not the input share equations. By contrast, the specification employed here introduces fixed effects into the input share equations in a logically consistent manner along with cross-equation constraints, thereby permitting unobserved network and related effects to influence input utilization.

[15]Hence, we assume a diagonal autocovariance matrix, with the diagonal elements for the share equation autoregressive parameters being equal. For further discussion, see Berndt and Savin [1975].

[16]These data were obtained at the state level and were then aggregated for each railroad according to the states through which each firm operates. Although such a method does not account for demand effects arising from interline traffic, any attempt to incorporate interlining would be ad hoc and would reduce the heterogeneity of the instruments. See Vellturo [1989] for a full discussion on the use of these variables and their construction.

[17]We deleted the linearly dependent M&S cost share equation from this system.

[18]The estimate of ρ in the cost function equation was 0.1535, with a t-statistic of 0.1160, while the estimate of the common ρ in the share equations was 0.1956, with a t-statistic of 0.1090.

[19]Specifically, point estimates of either the monotonicity or concavity conditions were violated at 119 of the 229 observations (52%).

[20]The point estimate of CC_{11} parameter was 0.6775, and dividing by the robust standard error yielded an asymptotic t-statistic of 2.0832.

[21]In estimation, the Atchison, Topeka and Santa Fe (ATSF) railroad was treated as the "base case" railroad; thus all other fixed effects estimates in Table 6 should be interpreted as differences from this base case railroad.

[22]The R^2 value for the variable cost function equation is 0.9988, while that for the labor, equipment and fuel share equations are 0.9395, 0.9721, and 0.8719.

[23]Note that this test is partial in the sense that, under the null hypothesis, the fixed effects of merged firms are not related to those of their pre-merged constituent firms, for this is not possible to do analytically. See Section IV for further discussion of this issue.

[24]The six mergers here involve CSX1 (1981), BNSL (1980), BNSYS (1982), NS (1982), UPSYS (1983) and CSX2 (1983).

[25] In these simulations, we assigned the input prices actually faced by merged firms post-merger to their constituent firms, and allocated capital stock and output levels of the merged firms in the same proportion as existed among constituent firms just prior to the merger.

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