Regional variations in cyclical employment

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Abstract. In this paper the author develops and tests a model of regional responses to national
business-cycles. The model divides cyclical decline in each state into two sectors: a basic sector and a nonbasic sector. The industrial mix, capital–labor ratio, age of capital stock, level of unemployment insurance benefits, labor shortage, and extent of labor-force unionization of a state are hypothesized to influence the response to national recessions by the economy of a state. Employment decline in the nonbasic sector of the economy of a state is a function of employment decline in basic industries and is transmitted through a short-run multiplier.

The model is tested on data from five post World War 2 recessions between 1950 and 1975. The findings indicate that industry mix at the two-digit Standard Industrial Code level explains 36% of the across-state variation in cyclical employment. The results also indicate that an old capital-stock, a nonunion labor force, and generous unemployment insurance benefits promote cyclical stability in state economies.

1 Introduction
National recessions are uneven in their spatial effects, with the older industrialized core regions of the United States of America experiencing more volatile business cycles than the rest of the country (Borts, 1960; Browne, 1978; Friedenberg and Bretzfelder, 1980; Gellner, 1974; Howland, 1979; 1981; Sum and Rush, 1975). Early work on regional cycles attributed these spatial variations in employment and income fluctuations to regional differences in industry mix. Isard (1957) hypothesized that regional “differences in the intensity and timing of regional cycles are explained in terms of differences in the sensitivity and responsiveness of particular industries. Cycles of a regional economy are simple composites of the cyclical movement of the economy’s industries appropriately weighted” (page 72). Empirical tests of the ‘industry mix’ hypothesis indicate that, although industry composition is important in explaining regional cycles, it alone does not explain cross-regional variations in employment and income (Borts, 1960; Browne, 1978).

The empirical work presented in this paper identifies institutional and structural factors that, in addition to industrial composition, influence the response of a region to national business cycles. The findings indicate that, if industry mix is held constant, the economies of states with old capital, a nonunion labor force, and generous unemployment insurance benefits tend to be cyclically stable.

Another strain of research on local cycles measures the responsiveness of local areas to national cycles by regressing a time series of local unemployment or employment against the comparable national series (Brechling, 1967; Thirwall, 1966). Recent evidence by Dunn (1982) demonstrates that regression coefficients estimated in this fashion are unstable, a result Dunn attributes to structural shifts in local and national economies. The results presented here suggest that temporal changes in local industry composition, age of capital, unionization of the labor force, and level of unemployment insurance benefits explain at least part of the instability noted by Dunn.

This study is carried out with a basic–nonbasic model of employment decline that incorporates the impact of a cyclical fluctuation in the basic industries of a
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The model is tested with state-level data from the United States of America for five recessions between 1950 and 1975.

2 Model

Each state economy is divided into basic-sector and nonbasic-sector employment. Basic activities, in this model, are defined as mining and manufacturing. Nonbasic activities include services, wholesale and retail trade, and government. This simplified division of basic–nonbasic activities may lead to an underestimation of nonbasic employment and the multiplier (μ), but does not bias the remaining coefficients. More complicated methods of allocating employment between basic and nonbasic industries, such as location quotients and minimum standard methods, would also result in biased values for μ (Greytak, 1969; Leigh, 1970; Richardson, 1979) and are not feasible given limitations in the data.

The dependent variable, the severity of the recession in a state, is equal to the weighted recessions in the basic sector and nonbasic sector of the state.

\[
E = E^b_T (E^m_T)^r + E^a_T (E^s_T)^r
\]

where
- \(E\) is the trend-adjusted severity of the recession;
- \(E^m\) is the trend-adjusted severity of the recession in the basic sector;
- \(E^s\) is the trend-adjusted severity of the recession in the nonbasic sector;
- \(E^m_T\) is the proportion of total employment in basic industries;
- \(E^s_T\) is the proportion of total employment in nonbasic industries;
- superscript \(j\) represents recessions 1, ..., 5;
- superscript \(r\) represents states 1, ..., 48;
- superscript \(t\) represents peak years 1, ..., 5.

The dependent variable (\(E\)) is the trend-adjusted percentage decline in state employment. It is detrended to eliminate distortions caused by regional differences in long-run growth. Without trend-adjustment, states experiencing strong secular growth would appear to experience milder recessions than states experiencing slow long-run growth (see Vernez et al, 1977; Howland, 1979). Whereas the National Bureau of Economic Research’s cycles are used to identify the national recessions, the peaks and troughs identified for each state are allowed to vary.

2.1 Basic sector

The severity of the recession in the basic-sector (\(E^m\)) of a state is divided into an industry mix component and a local component. The industry mix component is measured by weighting the trend-adjusted percentage change in employment in each national manufacturing industry by the importance of that industry in each state.

\[
E^m = \sum_{i=1}^{19} E^m_j \omega^i \cdot \omega^i
\]

where the notation is the same as before, and
- \(E\) is the industry mix component;
- \(E^m_j\) is the trend-adjusted severity of the recession in national industry \(i\);
- \(\omega^i\) is the proportion of employment in industry \(i\);
- superscript \(i\) denotes manufacturing industry at the two-digit Standard Industrial Code (SIC) level.
The remaining variation in basic-sector employment is explained by a local component—including five state-specific factors. The first two factors are the capital–labor ratio and the age of manufacturing capital, which are proposed to explain cross-state differences in regional output. The final three variables propose reasons for cross-state differences in layoff practices. These variables are the proportion of the labor force in unions in each state, the level of unemployment insurance benefits, and the existence of a labor shortage or surplus during the upswing of the cycle. Each of the five hypotheses are explained below.

2.1.1 Capital–labor ratio. The lower the capital–labor ratio of a state, the more severe its expected recession. The hypothesis applies both to multiplant firms with branches located across state boundaries and to single-plant firms.

During periods of cyclical downturn, managers of profit-maximizing multiplant firms should, ceterus paribus, allocate production cutbacks disproportionately with plants having high variable cost bearing a larger burden of economic slowdown than the plants having high fixed cost. The reason is that the cost of idle fixed inputs is borne entirely by the firm, whereas the cost of idle variable inputs is not, or is only partially, assumed by the firm. Labor is a major variable cost, whereas capital is a major fixed cost. Thus, losses to the firm are minimized when labor-intensive plants are made idle, workers are laid off, and production is shifted to capital-intensive plants. As a consequence, it is predicted that, during economic downturns, firms and, in the aggregate, states, with low capital–labor ratios will experience more severe reductions in aggregate output and therefore greater cyclical unemployment than their counterparts with high capital–labor ratios\(^1\). This hypothesis relies on evidence from Feldstein (1976), McLure (1977), and Vickery (1979). All three researchers found that with the current unemployment insurance system, firms do not bear the full cost of layoffs.

Differences in capital–labor ratios should also influence the bankruptcy or closure rates for single-plant firms in a competitive industry. Labor-intensive firms pay a higher proportion of total costs to variable factors than do capital-intensive firms. As average revenues (ARs) fall during the downswing of the cycle, the point where AR falls below average variable costs (AVC) will be earlier, ceterus paribus, for the higher-AVC firm than for the low-AVC firm. Thus labor-intensive, high-AVC firms, should shut down sooner than capital-intensive firms. The argument is not that total profits or losses will be greater in the labor-intensive or capital-intensive plant, rather that there will be different points when \( AVC = AR \), different rates of firm closures, and consequently more severe cyclical unemployment in regions where labor-intensive plants are concentrated. This analysis applies to the case of a competitive market with flexible prices as well as to the case where AVCs rise because of a reduction in output.

To illustrate the cross-regional variations in production functions within two-digit SIC level industries, the capital–labor ratios for the South, Northcentral, and Northeast were 12.0, 8.8, and 9.0, respectively, for textile manufacturing in 1972. The values for machinery manufacturing were 7.9, 7.6, 10.4, and 12.0 for the West, South, Northcentral, and Northeast regions, respectively, in 1972.

Because capital and skilled labor are complements in the production process, owners of capital-intensive firms may be reluctant to reduce output through layoffs because of the high cost of replacing skilled workers during the recovery. This effect would reinforce a negative sign on the capital–labor coefficient. Skill-level

\(^1\) It is possible that instead of plants with low capital–labor ratios being cyclically sensitive, plants in cyclically sensitive states produce with labor-intensive technologies, so as to allow flexibility during swings of the business cycle.
data, by state, for peak years between 1950 and 1978 are not available; therefore, the effect of skill levels on the severity of recessions could not be controlled. In addition, it is worth mentioning that capital-intensive operations are energy-using. Thus in the post-1973 period, these firms may have relatively high average costs.

2.1.2 Age of capital. The age of capital is also expected to influence the cyclical behavior of the economy of a state. The theory applies to profit-maximizing multi-plant and single-plant firms that respond to falling demand either by cutting prices or by maintaining prices and reducing output. The argument is as follows.

Plants with a capital stock of high average age are expected to experience high average costs for three reasons. First, old capital is less appropriate than new capital for current relative prices of land, labor, and other factors of production. Therefore, the old-capital plant is relatively costly to operate. Second, technological changes render old plants less cost-effective than new facilities. For example, multi-storey plants are less efficient because of the introduction of the assembly line. Last, old plants are less likely to be located in cost-minimizing locations than are new plants. Sources of raw materials and transportation modes shift with time. Moreover, because of technological changes in capital equipment, input mixes change and old locations are unlikely to minimize the transport costs of the new input mix. For any or all of these reasons, old-capital plants are expected to have high average costs.

For the multiplant firm, where prices are flexible and any level of output can be sold at the lower price, the manager continues producing at the old-capital plant as long as the price is greater than AVC. When prices fall below AVC, the old-capital plant will be closed. These closures will occur sooner in old-capital plants than in new-capital plants. When these closures are observed at the aggregate or macroeconomic level, the region with the capital stock of higher average age will experience greater cyclical unemployment than the region with the newer capital stock. A similar line of reasoning can be made for single-plant firms.

When prices are rigid and output levels are reduced in response to oversized inventories, reductions in output will also be concentrated in old-capital plants of multiplant firms. If output is cut because of falling demand for the products produced by firms, profits are maximized if cutbacks are made in the low-profit, high-average-cost plants. Layoffs, therefore, will be more severe in regions with relatively high proportion of old capital. This argument holds for oligopolists as well as for firms in more competitive industries. It is possible that average costs in the old-capital plant could fall below average costs in the new-capital plant as output levels are reduced. If this occurs, some reduction in output will occur in the new-capital plant. However, the old-capital plant will still experience a disproportionate share of cutbacks by the firm.

The assumption that new-capital stock is more appropriate for current relative prices should, however, be stated with some qualification. Relative energy prices fell slowly during the postwar period, 1945–1973, encouraging a transition toward energy-using capital. The well-known events of late 1973 led to a reversal of the energy-price trend and relative energy prices have increased. New energy-intensive capital may now be less efficient than older energy-saving plant and equipment, leading to higher marginal costs for the new-capital firm or plant. This particular change in relative prices would only affect the results of this study for the 1973–1975 recession. However, the possibility of other reversals in relative price trends necessitates a qualification of the hypothesis.

Variaya and Wiseman (1977) have suggested that an old capital-stock may lead to more severe regional recessions because the retirement of obsolete capital is
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concentrated in regions where the average age of capital is higher. During the expansionary phase of the cycle, scheduled retirements may be postponed, either because the revenues from running the old capital are temporarily higher than the salvage value of the land, labor, and capital; or because orders from regular or new customers must be met. With the end of the expansion, the delayed retirements combined with the regularly scheduled retirements are bunched together, creating the appearance of a more volatile cycle.

Additional evidence suggests that, contrary to the above hypotheses, the relationship between age of capital and the severity of regional cycles may be negative rather than positive. Rather than measuring efficiency of production, the age of capital stock may measure the average age of firms in the state. More clearly, states with a low average age of capital may be fast-growth states with a high proportion of small, new, and dynamic firms. New small firms are more likely to borrow credit to the limit, to make high-risk decisions, to be dependent on external funds, and therefore to be more susceptible to bankruptcy than their larger older counterparts. Birch (1979) found that small service and manufacturing firms are more likely to go bankrupt during recessions than are medium or large firms. He also found that the total number of jobs lost is greater during the downswing of the cycle in small firms than in medium or large firms. These findings suggest that states with a low average age of capital may be more cyclically sensitive than states with an older capital-stock, because the former have a high proportion of small, new, and cyclically vulnerable firms. Unfortunately data on the behavior of new firms are unavailable. Thus, although most authors hypothesize a positive relationship between the age of capital stock of a region and the severity of recessions, there is some reason to suspect the relationship may be negative.

The second set of three hypotheses to be explored in this model are factors proposed to explain cross-regional differences in layoff practices during a recession. The hypotheses suggest that holding industry mix constant, whether there is a labor surplus or shortage, the magnitude of unemployment insurance benefits (UI), and unionization of the work force will influence the cyclical employment of a region.

2.1.3 Labor surplus and labor. First, employers in labor-surplus markets may expect low labor-search costs during the recovery and, therefore, readily lay off workers during the downturn. Comparable firms in the labor-short states may anticipate difficulties in rehiring and, therefore, find it cheaper in the long run to hoard workers. Using the annual peak-level unemployment rate as a proxy, the greater the unemployment rate, the greater the expected recession. Thirwall (1966) found, using data from Great Britain, that regions experiencing the greatest cycle sensitivity were those with unemployment rates persistently above the national average.

2.1.4 Unemployment insurance benefits. Second, regions with greater UI relative to wages are expected to experience more severe regional recessions. The greater the UI in a state, in relation to wages, the more likely workers are to wait out the recession without looking for and taking another job. Employers, therefore, may be inclined to lay off workers, expecting them to be available for rehiring at a later date. Also, employees with some bargaining power are more likely to accept layoffs in high-UI states than in low-UI states. In low-UI states, workers may prefer wage or hour reductions to layoffs.

Hamermesh (1972) and Feldstein (1978) found that the level of UI has a positive effect on temporary layoffs at the firm level. The results of work by Holen (1977), Classen (1977), and Welch (1977), indicating that higher unemployment benefits lengthen the duration of unemployment, are consistent with the argument that rehiring is easier in high-UI states than in low-UI states.
A contradictory hypothesis by Welch (1977) suggests that an experience-rating system, even one that is not fully experience-rated, will raise the cost of laying off workers and consequently reduce layoffs. Since the payroll tax is greater in high-UI states than in low-UI states, the incentive for firms to hoard workers throughout the recession is stronger in states where UI is generous.

A final argument for greater stability in high-UI states than in low-UI states is that UI acts as an automatic stabilizer. The greater the benefits, the more stable the income level of a state, and the smaller the effects of the recession on the nonbasic sector.

2.1.5 Unionization of the labor force. Third, it is hypothesized that cross-regional differences in layoff practices arise from cross-regional differences in union strength. Feldstein (1978) and Medoff (1979) found evidence to support the hypothesis that workers in unionized firms have significantly higher probabilities of being laid off than workers in similar nonunionized firms. When demand for labor falls, management have several options for reducing their work force; to leave positions vacated by quits unfilled, to reduce or slow the growth in real wages, to reduce hours, or to increase layoffs.

Adjustments through unreplaced quits are less of an option for the unionized firm than for the nonunionized firm. The reason is that the quit rate in union firms is relatively low (see Freeman, 1978; Johnson, 1978).

A second option for labor adjustments is a reduction in wages. Empirical evidence by Hamermesh (1970) and Lewis (1978) suggests that real wages in the union sector are less sensitive to changes in the unemployment rate than are wages in the non-union sector; a finding that suggests that unionized establishments are unlikely to respond to falling labor-demand by reducing wages. With lower quit rates and less ability to reduce wages, union firms must make use either of layoffs or of work sharing.

Work sharing is likely to be the preferred strategy of the younger more recently hired workers. With work sharing, the marginal worker bears only part of the cost of the cutback whereas, with layoffs, the recently hired or marginal worker bears the total cost. But, the older workers prefer cutbacks to take the form of layoffs. Under a policy favoring layoffs, senior workers are likely to retain their jobs, and therefore incur no or little cost.

Because in nonunionized firms the marginal worker preference is transmitted to management, it is likely that cutbacks in such firms will take the form of work sharing and cuts in wages. In unionized firms where the demands of the average and more senior workers predominate, layoffs will be more likely to prevail (Medoff, 1979).

An additional hypothesized reason for the positive relationship between unionization and layoffs is that managers of unionized firms may find a policy favoring layoffs acceptable because they anticipate low rehiring costs during the recovery. Laid-off union workers are not likely to abandon a union job. Rather, they will collect unemployment benefits and wait to be recalled. This ensures the firm a ready pool of workers to draw from during the upswing, making firms less reluctant to lay off workers during the downturn. Additional evidence by Freeman (1978) has shown that years of tenure with an employer are positively correlated with unionization, a result consistent with the argument that workers are reluctant to relinquish a union job.

Union workers tend to be skilled. Since employers are reluctant to lay off skilled employees, the impact of unionization on the severity of regional recession will be muted.
2.1.6 Summary. To summarize, the recession in the basic sector of a state is:

\[ \hat{E}_m = \hat{E} + \hat{E}_L, \]  

where the variables are the same as before, except that:

\( \hat{E}_L \) is the local component of the recession in the basic sector of a state.

The local component (\( \hat{E}_L \)) includes the five economic and institutional factors described in sections 2.1.1 to 2.1.5 and is equivalent to:

\[ \hat{E}_L = \hat{E}(\beta_1R_{CL} + \beta_2A + \beta_3u + \beta_4U + \beta_5R_{UW}); \]  

where the \( \beta \)s are parameters to be estimated and the variables are the same as previously except that

\[ A tr = \frac{(I^{t} + I^{(t-1)})}{K^{t}}; \]

\( R_{CL} \) is the capital-labor ratio;

\( I \) is the total investment in fixed plant and equipment;

\( K \) is the value of the capital stock;

\( u \) is the percentage of the labor force belonging to unions;

\( U \) is the peak-year unemployment rate;

\( R_{UW} \) is the ratio of weekly unemployment insurance benefits to the average weekly wage.

The local component [equation (4)] includes the industry mix variable, \( \hat{E} \). This formulation gives more influence to the observations from states whose economies are cyclically sensitive because of industry mix. A small value for \( \hat{E} \) indicates that a state is comprised of cyclically stable industries, and consequently output and layoff adjustments should be small. In these states, capital-labor ratios, age of the capital stock, unionization, etc are not expected to have much influence on cyclical employment and are given less weight in the estimation of the \( \beta \) coefficients.

To summarize, it is proposed that within the basic sector of a state economy the severity of the actual recession deviates from the expected recession based on industry mix for five reasons. The capital-labor ratio, the age of the capital stock in the state, the extent to which its labor force is unionized, the existence of a labor shortage or surplus in peak years, and the level of unemployment insurance benefits vary across states and may influence the severity of state recessions. A final factor expected to influence the severity of the regional recessions is the difference in the decline in employment in the nonbasic activities of a state.

2.2 Nonbasic sector

The severity of the decline in nonbasic employment is assumed to be explained both by the severity of the employment decline in the basic sector because of industry composition and by the short-run multiplier (\( \mu \)). The multiplier is formulated as follows:

\[ \hat{E}^{tr} \mu = \left( \frac{E_m}{E_T} \right)^{rt} \hat{E}^{tr}. \]  

The parameter \( \mu \) is a short-run multiplier whose value should range between 0 and a long-run multiplier, \( (E_m/E_T)^{-1} \); \( \mu \) approaches the long-run multiplier the more prolonged the recession.

\( \hat{E}^{tr} = \mu(E_m/E_T)^{rt} \hat{E}^{tr} \) with \( \hat{E}_m \) estimated as shown in equations (2), (3), and (4). Coefficients in this model could not be estimated, however, because of the failure of the nonlinear maximum likelihood functions to converge for three of the five recession-specific equations. The results for the combined semipooled cross-section time-series equation of this model were similar to those shown in table 2.
The parameter $\mu$ is expected to be of small magnitude because it measures the effect of a temporary change in manufacturing employment on residentiary employment. Where employers expect the fluctuation in manufacturing to be temporary, they are less likely to cut output and lay off residentiary sector workers. Moreover, workers in the manufacturing sector are more likely to draw from savings and maintain current levels of demand for residentiary services when the downturn is expected to be shortlived.

2.3 Econometric model

The final equation, derived by substituting equations (2), (3), (4), and (5) into equation (1) is equal to:

$$
\hat{E}^{ir}_r = \left(\frac{E_m}{E_T}\right)^{rt} \left[ \hat{E}^{ir}_r \left( \beta_1 R_{CL} + \beta_2 A^{rt} + \beta_3 u^{rt} + \beta_4 U^{rt} + \beta_5 R_{UW} \right) \right] + \mu \left( \frac{E_s}{E_T} \right)^{rt} \left( \frac{E_m}{E_T} \right)^{rt} \hat{E}^{ir}_r.
$$

By use of ordinary least squares, equation (6) is estimated for each recession between 1950 and 1975. These recessions, as defined by the National Bureau of Economic Research, occurred in 1953-1954, 1957-1958, 1960-1961, 1969-1970, and 1973-1975. The model is also estimated with data pooled from all five recessions. The data sources are given in the appendix.

3 Results

In the five recession-specific equations in table 1, none of the six coefficients is consistently statistically different from 0. The strong results from the pooled equation, at the bottom of table 1 suggest that the poor recession-specific results arise from insufficient degrees of freedom and small variances in the independent variables (Rao and Miller, 1971). In light of the weak recession-specific results and the strong pooled results, the remainder of the discussion will focus on the latter.

An $F$-ratio was calculated to determine whether pooling of the data provides more precise estimates of the coefficients than can be estimated with nonpooled data (Maddala, 1977; Judge et al, 1980, chapter 8). The $F$-ratio for the pooled versus nonpooled equations in table 1 leads to the rejection of the stability hypothesis. There are significant differences in at least some of the coefficients across recessions.

Using dummies for each variable for each recession, equation (6) was reestimated. The stable coefficients were then combined and the unstable coefficients allowed to vary. The parameters for unionization and unemployment insurance benefits, and the multiplier were stable across all recessions and therefore pooled. The coefficients on the age of capital stock varied depending on the severity of the recession, as did the coefficients on the capital-labor ratios. The unemployment rate was stable for the last five recessions, but not the first. With this semipooled equation, the $F$-ratio was recalculated. The new $F$-ratio was 1.57, whereas the critical value is approximately 1.88 at the 1% level. Using the semipooled equation, we, therefore, do not reject the hypothesis of stability in the pooled coefficients.

In table 2 and in the rest of the paper, the superscripts on the parameters represent the recessions for which the coefficients were estimated. For example, $R_{CL}$ specifies that this coefficient on the capital–labor ratio was calculated with the data from the 1953–1954 recession.

As is frequently the case with cross-section data on states, there is evidence of heteroscedasticity in the error terms of the semipooled equation. The test for heteroscedasticity suggested by Glejser (1969) indicates that the error term is positively and significantly associated with the state population\(^3\).

\(^3\) The error term is given by $|e| = \beta(p)^{1/2}$, where $p$ is the population of the state and $\beta = 2546$, with a $t$-statistic of 9.7.
To make the necessary corrections, the semipooled equations were weighted by the square root of the population of each state. The corrected coefficients are displayed in the second column of Table 2. The results are discussed with reference to the corrected equation in Table 2.

The results for the capital-labor ratio variable ($R_{CL}$) indicate support for the hypothesis of a negative relationship between cyclical sensitivity and the capital-labor ratio of a state. The negative relationship hypothesized above appears to hold only when the severity of the national recession is relatively high in terms of real gross national product loss and is only statistically significant in the 1953-1954 recession estimation.

The age of capital stock ($A$) coefficients are estimated with a mixed degree of pooling. The coefficients for recessions 1953-1954, 1957-1958, and 1973-1975 are stable, consequently data from these periods are pooled; and the coefficients for recessions 1960-1961 and 1969-1970 are similar, so these two data sets are pooled. The coefficient on the age of capital stock is statistically different from 0 at the 0.01 level for the more severe 1953-1954, 1957-1958, and 1973-1975 recessions. The coefficient is insignificant for the milder 1960-1961 and 1969-1970 recessions.

In contrast to the negative relationship hypothesized by most researchers, the coefficient $A_{1.2.5}$ is positive. A positive parameter suggests that states with a newer manufacturing capital-stock experience more-severe recessions than do states with a

| Table 1. Final results for the five recession-specific equations and the pool equations. [The results are calculated from equation (6) in the text; the parameters are also described in the text; the statistics in ( ) are standard errors, and those in [ ] are $t$-statistics.] |
|---------------------------------------------------------------|
| $R_{CL}$ | $A$ | $u$ | $U$ | $R_{UW}$ | $\mu$ | $F$-ratio | $SSR^a$ | $n^b$ |
|---------|-----|-----|-----|--------|------|----------|--------|-----|
| 1953-1954 recession | -0.11 | 18.92 | 0.01 | 0.27 | -15.19 | 0.97 | 3.41 | 39.77 | 33 |
| (0.360) | (7.411) | (0.040) | (0.220) | (6.606) | (0.620) | | | |
| [0.29] | [2.55] | [0.28] | [1.21] | [-2.30] | [1.56] | | | |
| 1957-1958 recession | 0.21 | -5.97 | 0.02 | -0.15 | -7.24 | 1.92 | 0.79 | 22.34 | 35 |
| (0.224) | (6.817) | (0.022) | (0.189) | (3.312) | (0.449) | | | |
| [0.93] | [0.88] | [1.04] | [-0.77] | [-2.219] | [4.29] | | | |
| 1960-1961 recession | 0.03 | -4.13 | 0.03 | -0.32 | 0.71 | 1.07 | 1.94 | 18.3 | 37 |
| (0.136) | (11.60) | (0.029) | (0.234) | (4.653) | (0.720) | | | |
| [0.24] | [0.36] | [1.08] | [-1.37] | [0.15] | [1.48] | | | |
| 1969-1970 recession | -0.00 | -19.25 | 0.03 | -0.68 | 15.08 | -0.26 | 6.60 | 12.69 | 40 |
| (0.087) | (8.673) | (0.037) | (0.463) | (7.116) | (0.417) | | | |
| [0.04] | [2.23] | [0.71] | [-1.47] | [2.12] | [-0.63] | | | |
| 1973-1975 recession | -0.06 | 11.09 | 0.01 | -0.12 | -4.85 | 0.71 | 10.472 | 18.88 | 47 |
| (0.044) | (5.402) | (0.022) | (0.168) | (3.680) | (0.279) | | | |
| [1.47] | [2.05] | [0.61] | [-0.758] | [-1.32] | [2.55] | | | |
| All recessions | -0.08 | 6.47 | 0.03 | -0.01 | -5.91 | 0.89 | 29.06 | 144.29 | 192 |
| (0.033) | (1.968) | (0.010) | (0.084) | (1.859) | (0.199) | | | |
| [2.32] | [3.29] | [2.59] | [0.13] | [-3.18] | [4.49] | | | |

$^a$ SSR is sum of squared residuals.

$^b$ $n$ is the number of states in the calculation; a number of states had to be excluded from the analysis when data on the independent variables were missing. The excluded observations tended to be the smaller less industrialized states, where data were suppressed to keep information on individual firms from being revealed.
relatively old stock of capital. It is worth noting here that a high value for the age of capital variable signifies a new capital-stock.

One explanation for the positive sign on age of capital is that rather than measuring ‘efficiency’ of production, the age variable measures the average age and size of firms in the industry. As stated above, a positive relationship between the size of a firm and its resiliency during recessions has been argued by Birch (1979). Regions with new capital may have a high proportion of new small firms that have a high probability of going bankrupt during recessions. To find out whether the age of the capital stock of a state is associated with the average size of its firms and the number of its bankruptcies, the age variable is correlated with the percentage of manufacturing firms with twenty or more employees in a state and with the percentage of manufacturing firms completing bankruptcy proceedings in each state in 1973–1975. The simple correlation between the age-of-capital-stock variable and bankruptcy is 0.49. That is, states with new capital experience higher rates of bankruptcy during recessions. The simple correlation between age of capital and the size of firms is −0.52, indicating states with new capital tend to have smaller firms.

The insignificance of the coefficient on the age-of-capital-stock variables for the pooled 1960–1961 and 1969–1970 recessions can be explained by the mildness of these two recessions. If high rates of bankruptcy explain the negative relationship between age of capital stock and severity of state recessions (or the positive sign between the age-of-capital-stock variable and severity), then it is likely that most firms,

Table 2. Uncorrected semipooled results and semipooled results corrected for heteroscedasticity. [The results are calculated from equation (6) in the text; the parameters are also described in the text; the statistics in ( ) are standard errors, and those in [ ] are t-statistics.]

| Parameter | Results | Parameter | Results |
|-----------|---------|-----------|---------|
|           | uncorrected | corrected | uncorrected | corrected |
| $R_{CL}$  | −0.256  | −0.45     | −0.58    | 1.33    |
|           | (0.144) | (0.152)   | (4.463)  | (4.28)  |
|           | [−1.78] | [−2.99]   | [−0.13]  | [0.31]  |
| $R_{CL}$  | −0.18   | −0.167    | 0.03     | 0.03    |
|           | (0.092) | (0.087)   | (0.0119) | (0.0100)|
|           | [−2.00] | [−1.94]   | [2.53]   | [2.70]  |
| $R_{CL}$  | 0.158   | 0.24      | 0.05     | 0.09    |
|           | (0.08)  | (0.082)   | (0.107)  | (0.139) |
|           | [1.74]  | [2.99]    | [0.45]   | [0.68]  |
| $R_{CL}$  | −0.04   | 0.01      | −0.14    | −0.24   |
|           | (0.077) | (0.069)   | (0.107)  | (0.092) |
|           | [−0.48] | [0.19]    | [−1.33]  | [−2.59] |
| $R_{CL}$  | −0.08   | −0.05     | −5.64    | −4.73   |
|           | (0.039) | (0.037)   | (1.905)  | (1.884) |
|           | [−2.15] | [−1.48]   | [2.96]   | [2.51]  |
| $A^{1,2,5}$ | 9.53   | 13.60     | 0.94     | 0.40    |
|           | (2.906) | (2.80)    | (2.021)  | (1.75)  |
|           | [3.28]  | [4.85]    | [4.66]   | [2.27]  |

a The superscripts indicate the recession data used to calculate the parameter: 1 is the 1953–1954 recession; 2 is 1957–1958; 3 is 1960–1961; 4 is 1969–1970; 5 is 1973–1975.

b The correction uses weighted least squares: weighted by $p^{-1/2}$ where $p$ is the state population.

c SSR is sum of squared residuals.

d $n$ is the number of states in the calculation.
firms, independent of age and size, weathered the milder 1960–1961 and 1969–1970 downturns.

The insignificance of the 1960–1961 and 1969–1970 age-of-capital-stock coefficient suggests that, during a relatively mild recession, firms independent of size, age, and labor-intensiveness, are likely to survive through such mechanisms as extensions of credit or savings. Firms faced with reductions in demand are able to reduce output through reductions in hours worked and increases in inventories without layoffs. Recent work by Clark (1983) supports the argument that firms tend to adjust hours worked before adjusting employment. These factors may explain why the age of capital stock as well as capital–labor ratios have little influence on the severity of cycles of state economies during moderate national recessions.

The coefficients on unionization \((u)\) are extremely stable across all recessions, and all five samples were pooled to estimate one coefficient. As expected, the coefficient is positive and statistically different from 0 at the 0.01 level. The coefficient of 0.03 suggests that an increase in the proportion of the labor force in the state belonging to unions will lead to greater amplitude in the employment cycle in the state. For example, unionization explains a 5.7% loss in basic-sector employment in Michigan [5.8 (38.4 x 0.03)] and a 2% loss in basic-sector employment in North Carolina [8.7 (7.5 x 0.03)] during the 1973–1975 recession. Michigan is the most heavily unionized state in the United States of America, whereas North Carolina is the least unionized state. In short, the evidence does indicate that at least part of the cross-state variation in employment cycles can be explained by the degree to which the labor force is unionized.

The coefficients on the peak-level unemployment rate \((U)\) are stable for the four recessions between 1956 and 1975. Consequently, this coefficient is estimated with data from all four recessions. The coefficient for the 1953–1954 data is estimated separately. Contrary to the hypothesis and evidence from Great Britain (Thirwall, 1966), the coefficient on \(U^{2,3,4,5}\) is negative, \(-0.24\), and statistically significant at the 0.025 level. The coefficient on \(U^1\) is not statistically different from 0.

The negative statistically significant sign on \(U^{2,3,4,5}\) is not readily explained by economic theory, but may be explained by misspecification of the model. The peak-level unemployment rate is hypothesized to explain the severity of the subsequent recession. It is, however, possible that the unemployment rate is not independent of the severity of the previous recession. For example, a state with cyclically volatile cycle may not fully recover from the last downturn before the peak-level unemployment rate is registered. If the next recession in the state is relatively mild, we might get a high rate of unemployment with a mild state-recession. The variable \(U\) was dropped from the equation to determine whether this misspecification was distorting other coefficients. Eliminating \(U\) made no noteworthy changes in the magnitude or statistical significance of the remaining parameters.

The parameter on the ratio of average weekly unemployment insurance benefits to weekly wages \((R_{uw})\) is estimated with the pooled data. As shown in table 2 the pooled coefficient is negative and statistically significant at the 0.02 level. The coefficient of \(-4.7\) implies that the current unemployment insurance program is stabilizing. The greater \(R_{uw}\) benefits, the more stable the income level of a state and the smaller the effects of the recession on the residentiary sector. Using the Michigan example, the 1972 ratio of unemployment insurance benefits to weekly wages in Michigan of 0.319, reduced cyclical unemployment in that state by 8.8% [5.8 \((-4.73 \times 0.319\)].

The short-run multiplier, \(\mu\), is an average value for all states for all five recessions and it measures the impact of the severity of the basic-sector recession on the residentiary sector. The model estimates \(\mu\) to be approximately 0.4 and the
The coefficient is statistically significant at the 5% level. This result is consistent with theory as well as with empirical evidence.

What proportion of the variation in recessions in states is explained by the model? To permit the calculation of an adjusted $R^2$-value, an equation including the constant ($\alpha$) and the expected recession:

$$\hat{E}^{fr} = \alpha + \beta \hat{E}^{fr},$$  

and an equation including a constant, the expected recession, and the local component:

$$\hat{E}^{fr} = \alpha + \beta_0 \hat{E}^{fr} + \hat{E}^{fr} (\beta_1 R_{CL} + \beta_2 U^{rt} + \beta_3 U^{rt} + \beta_4 R_{Iw}^t),$$

are estimated to determine the extent to which the local component and the resideniary component, contribute to the explanatory power of the model. The variables are the same as defined above.

The results indicate that industry mix component alone explains 36% of the variation in state cycles. The inclusion of the local component explains an additional 14% of the variation, whereas the contribution of the nonbasic sector in the form of the complete model contributes little to the $R^2$-value. The model, therefore, explains 50% of the variation in the dependent variable. The unexplained variation may be explained by heterogeneity within two-digit SIC categories.

To estimate how much variation in the dependent variable can be explained by heterogeneity within two-digit SIC categories of the industry mix variable, the severity of the recession in the machinery and textile equations are regressed, using ordinary least squares, against a constant and a three-digit SIC industrial composition variable.

$$\hat{E}^{fr} = \sum_{k=1}^{9} (\hat{E}^{fr} \omega^k),$$

The variables are the same as before except:

$k$ is the industrial detail at the three-digit SIC level.

The results suggest that the degree of heterogeneity depends upon the industry. For example, there appears to be a substantial degree of variation in SIC 35 (machinery manufacturing) explained by industry composition at the three-digit level. The machinery equation suggests that 38% of the cross-state variation in employment fluctuations in SIC 35 can be explained by industry composition at the three-digit SIC level. According to the findings from SIC 22, none of the cross-state variation in the textile recessions is explained by industry composition at the three-digit level. These results suggest that, if the industry mix or national component were calculated with a finer industry breakdown, the model would explain more of the regional variation in cyclical employment.

4 Conclusion
In this study cross-section time-series data are used to test the extent to which local economic and institutional variables influence local responses to national employment cycles. The findings suggest that 36% of the cross-state variation in employment cycles can be explained by cross-state differences in industry mix. Evidence was also found that there are cross-state differences in several economic and institutional variables that influence the severity of state business-cycles. If industry mix is held constant, the economies of states with an old capital-stock, that are nonunionized, and that distribute high unemployment insurance benefits tend to be cyclically stable.
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APPENDIX

Data sources
Severity of recession ($\hat{E}$) percentage deviation of seasonally adjusted data from a five-year moving average of the data at the trough of the recession, calculated from the Bureau of Labor Statistics monthly nonagricultural employment data (available on computer tape from the Bureau).

Industry composition weights ($\omega$) by state Census of Manufactures (USBC, 1954; 1958; 1963; 1967; 1972a) and Annual Survey of Manufactures (USBC, 1951–1976).

Value of capital stock by state estimated with the perpetual inventory method and base-year book values of depreciable assets from the “Special geographical supplement to 1962–64 data on book value of fixed assets and rental payments for buildings and equipment, 1972” Annual Survey of Manufactures (USBC, 1972b); an investment series taken from the Census of Manufactures (USBC, 1954; 1958; 1963; 1967; 1972a) and the Annual Survey of Manufactures (USBC, 1951–1976); and an implied depreciation rate calculated from two years of book values of depreciable assets and intervening years of investment.

Manufacturing employment by state Census of Manufactures (USBC, 1954; 1958; 1963; 1967; 1972a) and the Annual Survey of Manufactures (USBC, 1951–1976).

Age of capital stock ($A$) by state Census of Manufactures (USBC, 1954; 1958; 1963; 1967; 1972a) and Annual Survey of Manufactures (USBC, 1951–1976).

Percentage of the labor force belonging to unions, by state ($u$) extrapolations of data taken from Statistical Abstract (USBC, 1957; 1962); the Directory of National and International Labor Unions in the US, 1969 (USDL, 1970); and The Directory of National Labor Unions and Employee Associations, 1973 (USDL, 1974).

Ratio of average weekly unemployment insurance benefits to the average weekly wage ($R_{uw}$) Handbook of Unemployment Insurance, Financial Data, 1938–76 (USDL, 1978)

Peak-year unemployment rates by state ($U$) Manpower Report of the President (USDL, 1964; 1973).

Bankruptcy rates, by state US Courts data base, unpublished data.

Number of firms with twenty or more employees, by state Census of Manufacturers (USBC, 1972a).