

FLUCTUATIONS IN THE PACE OF LABOR REALLOCATION

STEVEN J. DAVIS*
University of Chicago

I. THREE HYPOTHESES ABOUT UNEMPLOYMENT-RATE FLUCTUATIONS

Familiar observations: The pattern and distribution of production opportunities change constantly in a modern economy. These changing opportunities spur a time-consuming, and otherwise costly, reallocation of specialized resources. In the case of labor, this reallocation process often manifests itself as unemployment.

Most readers will agree that these observations are quite pertinent to an economic explanation of the underlying rate of unemployment in a market economy. Upon turning to an explanation of unemployment-rate fluctuations, or the broad pattern of fluctuations in aggregate economic activity, I suspect that the same readers will attach small weight to fluctuations in the magnitude of specialized resource reallocation. The main reason that most economists assign a backseat status to the reallocation process when explaining aggregate economic fluctuations is set forth succinctly by Lucas (1977, p. 20):

A new technology, reducing costs of producing an old good or making possible the production of a new one, will draw resources into the good which benefits, and away from the production of other goods. Taste shifts in favor of the purchase of one good involve reduced expenditures on others. Moreover, in a complex modern economy, there will be a large number of such shifts in any given period, each small in importance relative to

*I thank David Altig, Karl Brunner, Robert Hall, John Haltiwanger, James Hamilton, John Huizinga, Kevin Murphy, Kathryn Shaw, Lawrence Summers, Robert Topel, Rob Vishny, and, especially, Walter Oi for helpful conversations and comments. Mark Rush kindly supplied some of the data used in this paper. The Graduate School of Business, University of Chicago provided financial support.

total output. There will be much "averaging out" of such effects across markets. [Italics and quotation marks appear in the original.]

A second reason for the backseat status is that a focus on the reallocation process seems to offer little insight into intriguing patterns of comovement between nominal and real variables.

Despite these arguments for a backseat status, some economists hold sharply different views on the role of the reallocation process in aggregate fluctuations. Black (1982) expresses one alternative view:

Both human and physical capital will be specialized, because specialization increases expected productivity. However, there are shocks to both tastes and technology in the form of unexpected shifts in the state variables in the economy. As a result, we find ourselves with a capital stock whose composition is different from the composition we would have chosen had we known in advance what the world would be like. The match between resources and wants is not perfect. ...Shocks that create a poor match, which will mean large shifts of human capital between sectors, cause unemployment and a decline in output.

Black proceeds to sketch a general equilibrium theory that treats fluctuations in the degree of mismatch between actual and desired resource allocation as a major driving force behind business-cycle fluctuations. Davis (1986a) and Hamilton (1986) consider simpler, but more fully articulated, models that exhibit unemployment-rate fluctuations driven by fluctuations in the degree of mismatch between the desired and actual allocation of labor resources. In these models the effects of allocative disturbances do not "average out," because (1) the reallocation of specialized labor is costly and time-consuming and (2) the arrival pattern of information about the desired labor-force allocation is uneven over time. The view that a large fraction of unemployment-rate fluctuations is driven by an uneven arrival pattern of information about the desired labor-force allocation, I shall call the sectoral shifts hypothesis.

The purpose of this paper is to evaluate the sectoral shifts hypothesis and two competing hypotheses, to be introduced shortly. In evaluating these hypotheses, I draw on previous empirical findings as well as several new empirical findings reported in this paper. In some cases I offer new interpretations of earlier empirical findings. I focus on the

role, if any, that fluctuations in the pace of labor reallocation play in short-run unemployment-rate fluctuations. The interesting question--does an increase in labor market mismatch explain the increase in the average unemployment rate since 1974 relative to the preceding twenty-five years?--is not my main object of attention. For recent studies that bear on this and related questions, see Medoff (1983) and Topel (1986).

Before introducing the competing hypotheses, it is useful to recount Lilien's (1982) empirical finding and the provocative interpretation that he attributes to this finding. Using employment data, disaggregated by industrial sectors, Lilien constructed a time-series measure of the cross-sectoral dispersion in employment growth rates. He reasoned that relatively high values of the dispersion measure would accompany a rapid pace of sectoral labor reallocation and that relatively low values would accompany a milder pace of reallocation. He included this dispersion measure and its lag as regressors in a reduced-form unemployment-rate equation that also contains current and lagged values of a monetary disturbance measure. Using annual, postwar data for the United States, Lilien found a statistically and economically significant positive relationship between the dispersion measure and the unemployment rate. Based on this finding, Lilien (1982, p. 778) argues that "[a]s much as half of the variance of unemployment over the postwar period can be attributed to fluctuations of the natural rate brought about by the slow adjustment of labor to shifts of employment demand between sectors of the economy."

Lilien's interpretation is certainly consistent with his empirical finding, but it is open to question. Questions about his interpretation arise because the dispersion measure does not directly measure the quantity of labor reallocation. Hence, movements in the dispersion measure could reflect something besides variations in the pace of labor reallocation--for example, the normal sectoral pattern of temporary layoffs and recalls over

the business cycle.¹ Abraham and Katz (1986) describe two sets of conditions under which aggregate demand disturbances generate movements in the dispersion measure and a positive correlation between the dispersion measure and the unemployment rate. These two sets of conditions are: (1) a positive correlation between the change in and level of the unemployment rate, and either (2a) a negative correlation between sectors' trend growth rate and cyclical sensitivity, or (2b) sectors "differ in their cyclical sensitivities and labor force adjustment costs are asymmetric such that an increase in employment costs more than a decline of equal magnitude" (see, also, Weiss (1986)). Abraham and Katz show that conditions (1) and (2a) are satisfied in postwar United States data and suggest that condition (2b) may also be satisfied. Thus, the analysis by Abraham and Katz shows that Lilien's empirical finding is consistent with what I shall call the normal business-cycle hypothesis. The normal business-cycle hypothesis asserts, with respect to aggregate economic fluctuations, that the effects of allocative disturbances largely average out and that variations in the pace of labor reallocation are an unimportant aspect of unemployment rate fluctuations. Clearly, this hypothesis encompasses a wide class of very different models that share a lack of emphasis on the role of variations in the pace of labor reallocation. This class includes equilibrium and Keynesian monetary models of the business cycle as well as most nonmonetary models of the business cycle. The normal business-cycle hypothesis accords nicely with most formulations of the natural-rate hypothesis and the common practice of dividing the unemployment rate into a cyclical component, on the one hand, and structural and frictional components, on the other hand.

In contrast, the reallocation-timing hypothesis asserts that intertemporal substitution effects lead to a concentration of labor reallocation during recessions. According to the reallocation-timing

¹A further difficulty arises because the dispersion measure responds directly only to net intersectoral labor reallocation. Lilien's dispersion measure, and some of the measures used in this paper, are constructed from nonagricultural employment data disaggregated into a small number of industrial sectors. These measures could conceivably disguise or misrepresent dramatic fluctuations in the magnitude of interindustry labor flows or gross interindustry labor flows. Quantitatively, the size of the interindustry flows and the gross interindustry flows dwarfs the size of the net interindustry flows; see Leonard (1987). Hence, a Lilien-type dispersion measure is a useful device for measuring the effects of fluctuations in labor reallocation on unemployment only if it proxies for some broader concept of labor mobility. Section II presents evidence indicating that the pace of net labor reallocation across industries—as measured by Lilien's dispersion measure—exhibits a pattern of fluctuations similar to the pattern in the pace of labor reallocation across jobs.

hypothesis, the pace of labor reallocation and the accompanying structural and frictional unemployment fluctuate significantly over the business cycle. Why? The reallocation of specialized resources, like labor, involves foregone production due to lost work time (unemployment) and adjustment costs that take the form of foregone output. On average across sectors, the value of this foregone production fluctuates procyclically, rising in good times and falling in bad times. Thus, if labor mobility and turnover are substitutable over time, fluctuations in the average value of production induce fluctuations in the pace of labor reallocation. This phenomenon is illustrated in an explicit multisector model in my 1986(a) paper, where I call it the intertemporal substitution of labor mobility to emphasize its close parallels to the intertemporal substitution of leisure.²

A simple conceptual framework is useful for thinking about the sectoral shifts and reallocation-timing hypotheses. Consider a multisector economy with time-consuming labor mobility. At a point in time, there is some actual sectoral allocation of labor resources. At the same point in time, there is an implied target allocation to which the economy will eventually converge in the absence of further allocative disturbances. Of course, new allocative disturbances arrive over time, so that the target allocation continually moves away from (or toward) the actual allocation. Unemployment in this economy reflects the adjustment process by which labor mobility drives the actual allocation in the direction of the target allocation. Unemployment is an increasing function of the distance between the actual and target labor allocations, so that the arrival of new information about the desired labor-force allocation triggers an unemployment response by shifting the target allocation. The sectoral

²The very different model of Jovanovic (1987) also exhibits this phenomenon. Darby, Haltiwanger, and Plant (1985, p. 628) recognize the potential importance of this phenomenon, writing "Essentially, we are arguing that structural change in the economy will be greater during recessions. The basic idea is that the necessary reallocation of labor associated with changing tastes and technology is likely to be bunched together during recessions. This is because during these periods the value of production is relatively low, and therefore this becomes an optimal time to make changes that were eventually going to be made anyway."

Rogerson (1986) analyzes a different mechanism that affects the timing of labor reallocation. He shows how the timing of secular movements between sectors with differential trend productivity growth rates depends on the relative cyclical sensitivities of the two sectors. In his model workers discount future returns and costs, and they face a fixed cost of sector-switching that varies across workers. In this setting, if the shrinking sector is more (less) cyclically sensitive, then the intersectoral flow of workers accelerates (decelerates) during recessions.

shifts hypothesis asserts that the arrival pattern over time of such information is sufficiently uneven as to induce substantial unemployment-rate fluctuations. The point of the reallocation-timing hypothesis is that, holding fixed the target allocation, fluctuations in the opportunity cost of labor mobility affect the timing of movement toward the target allocation. Similarly, in an economy with a target allocation that moves away from the actual allocation at a constant rate--most stationary equilibrium search models fit this description--fluctuations in the opportunity cost of labor mobility induce unemployment-rate fluctuations. More generally, in a world with a moving target, the unemployment rate is (1) an increasing function of the distance between the actual and target allocations and (2) a decreasing function of today's opportunity cost of mobility relative to the expected future opportunity cost of mobility. For analysis of an explicit dynamic equilibrium model that fits within this framework and illustrates these propositions, see Davis (1986a).

These remarks make it clear that the reallocation-timing hypothesis is also consistent with Lilien's empirical finding. Like the sectoral shifts hypothesis, the reallocation-timing hypothesis interprets Lilien's finding as evidence that large, short-run movements in the unemployment rate are associated with variations in the pace of labor reallocation. Unlike the sectoral shifts hypothesis, the reallocation-timing hypothesis asserts that aggregate--"average" is a more accurate characterization--disturbances largely determine the timing of recessions and expansions. Both the sectoral shifts and reallocation-timing hypotheses incorporate an explanation for why the effects of allocative disturbances might not "average out," but the explanations are quite different. Unlike the normal business-cycle hypothesis, the reallocation-timing hypothesis suggests that the severity of recessions can be strongly affected by the accumulated degree of mismatch between the desired and actual labor-force allocations. Also, whereas the normal business-cycle explanation for Lilien's finding relies on differences in sectors' cyclical sensitivities, these differences are inessential to the reallocation-timing explanation.

Since each of the three competing hypotheses can explain the positive correlation between dispersion measures and the unemployment rate, I assess these hypotheses by considering a number of corollary hypotheses or predictions. In considering these corollary hypotheses, several further differences among the competing hypotheses emerge. To sharpen the exposition, I frequently frame the discussion in terms of strong-form versions of the competing hypotheses, but it should be clear that weak-form

versions of the hypotheses are not mutually exclusive. Conceivably, elements of all three hypotheses enter into a full explanation for the positive correlation between dispersion measures and the unemployment rate.

Readers may appreciate a brief roadmap to remaining sections of the paper. Section II presents evidence that fluctuations in the pace of labor reallocation are an important element of fluctuations in the unemployment rate. This evidence emerges by considering the cyclic behavior of unemployment inflow and outflow rates, the labor-force participation rate, and temporary and permanent job separations. Section II also briefly considers the time-series relationship between job-vacancy measures and the pace of labor reallocation. Section III develops and tests two previously unexploited implications of the sectoral shifts hypothesis. The sectoral shifts hypothesis predicts a particular stage-of-the-business-cycle effect in the time-series relationship between the unemployment rate and the cross-sectoral dispersion measure. The sectoral shifts hypothesis also predicts that past patterns of labor reallocation affect current unemployment behavior. Section III introduces cross-sectoral covariance measures that index the current direction of labor reallocation relative to past directions. Using annual data over a 1924 to 1985 sample, section III presents strong evidence that past patterns of labor reallocation affected unemployment in the predicted way, particularly over the 1924 to 1946 period. Section IV tests the basic prediction of the reallocation-timing hypothesis that the pace of labor reallocation varies inversely with the cross-sectoral average value of foregone production. This prediction is confirmed using several different proxies for the value of foregone production. Section V summarizes the messages of section II-IV for the three competing hypotheses and draws conclusions.

II. CYCLIC VARIATIONS IN THE PACE OF LABOR REALLOCATION

A fundamental implication of the sectoral shifts and reallocation-timing hypotheses is that an increased pace of labor reallocation will accompany short-run increases in the unemployment rate. Given the difficulties surrounding the interpretation of Lilien-type dispersion measures, in this section of the paper I pursue an alternative device for gauging the pace of labor reallocation. Specifically, I consider the pattern of cyclic variation of flows into and out of the unemployment pool as an indicator of the pace of labor reallocation. An increase (decrease)

in the pace of labor reallocation implies coincident increases (decreases) in unemployment inflows and outflows. Thus, the sectoral shifts and reallocation-timing hypotheses predict that increases in unemployment inflows and outflows will accompany short-run increases in the unemployment rate. This prediction is the first matter I investigate in this section.

The second matter I consider is the empirical relationship between unemployment inflows and outflows and the proxy for the job vacancy rate used by Abraham and Katz (1986). My findings in this regard, and the accompanying discussion, lead me to reject the Abraham Katz interpretation of the negative time-series relationship between the vacancy rate proxy and Lilien's dispersion measure.

A. UNEMPLOYMENT INFLOWS AND OUTFLOWS

To construct unemployment inflow and outflow measures, I use monthly data from the Current Population Survey on the civilian labor force, the number unemployed, and the number unemployed less than five weeks. Let n_t denote the civilian labor force at time t , s_t denote the number of unemployed workers at time t , and s_t^{0-4} denote the number of workers unemployed less than five weeks at time t (more precisely, the number of workers unemployed 31 days or less). The monthly flow into unemployment, as a percentage of the labor force, equals the number of newly unemployed divided by the labor force:

$$\phi_t = \frac{s_t^{0-4}}{n_t} .$$

The monthly flow out of unemployment, as a percentage of the labor force, equals the difference between the number unemployed at $t-1$ and the number of previously unemployed who remain unemployed at t , divided by the labor force:

$$\omega_t = \frac{s_{t-1} - (s_t - s_t^{0-4})}{n_t} .$$

The unemployment rate is simply

$$UN_t = \frac{s_t}{n_t} ,$$

and, for small changes in the labor force, the month-to-month change in the unemployment rate approximates the difference between the inflow and

outflow rates:

$$\Delta UN_t \approx \frac{s_t - s_{t-1}}{n_t} = \phi_t - \omega_t.$$

This procedure for measuring unemployment inflows and outflows is used by Darby, Haltiwanger, and Plant (1985, 1986). Since Kaitz (1970), similar procedures have been widely used to analyze the distribution of unemployment-spell durations.

The inflow and outflow measures have several shortcomings. First, these discrete-time measures understate inflows and outflows to the extent that some individuals move into and out of unemployment between survey dates. This problem is minimized by the use of monthly data. Second, the calculation of ω_t uses s_{t-1} to estimate the total number of unemployed one month (31 days) ago, but the time span between surveys is either four or five weeks. Thus, s_{t-1} and s_t^{0-4} are not perfectly aligned at 31 days apart. Darby, Haltiwanger, and Plant (1985) show that, for reasonable values of the unemployment rate and unemployment exit probability, the bias introduced by this misalignment is negligible. Third, survey respondents' unemployment-spell durations are subject to a substantial amount of reporting error. Poterba and Summers (1984) use a month-to-month matched sample of survey respondents to investigate the consistency of individuals' responses from one month to the next. For our purposes, their most relevant findings pertain to the reported spell durations of "new entrants" into unemployment—those who were classified as employed or not in the labor force in month one and unemployed in month two. About four and one-half weeks elapsed between month one and month two in the Poterba and Summers sample, but roughly one-quarter of all new entrants into unemployment in month two reported spell durations of five weeks or more, and about eight percent of these new entrants reported spell durations of twenty-five weeks or more. By considering a three-month matched sample, Poterba and Summers provide evidence that many of these long reported spell durations arise because individuals typically lump intermittent short spells into one long spell when responding to questions about the duration of an ongoing unemployment spell. The magnitude of response error argues for a cautious stance in the interpretation of the flow measures. These measures may misstate the average level and the amplitude of fluctuations in the "true" flows through the pool. However, any prominent cyclic patterns in the actual flows will likely emerge in the measured flows. I concentrate on these cyclic patterns.

Two other factors complicate the interpretation of cyclic variation in the flow rates. First the ϕ and ω measures do not distinguish between flows to or from employment and flows to or from not in the labor force. For example, an increase in the outflow rate, ω , can reflect either a more rapid pace of job creation, or a discouraged-worker effect whereby more individuals leave the labor force. To address this difficulty, I briefly consider the cyclic behavior of the labor-force participation rate, but a fully satisfactory treatment clearly requires an analysis of the Current Population Survey gross flow data on transitions between employment, unemployment, and not in the labor force. I hope to pursue this avenue in future research. Abowd and Zellner (1985), Poterba and Summers (1986), and others have recently developed procedures designed to adjust the gross flow data for reporting errors and other problems. Second, the outflow measure, ω , does not distinguish between workers who return to their previous job and workers who switch to a new job. This shortcoming makes it difficult to assess the relative contributions of permanent and temporary separations to fluctuations in the unemployment flows. To address this shortcoming of the flow measures, I consider some direct evidence on the importance of permanent and temporary separations.

Bearing these shortcomings in mind, Figure 1 plots quarterly averages of the monthly flow measures and unemployment rate for the 1948:1 to 1986:2 time-period.³ In line with the approximation, changes in the unemployment rate are roughly three times the difference between the average monthly inflow rate and the average monthly outflow rate.

Two patterns emerge clearly from Figure 1. Short-run increases (decreases) in unemployment are usually accompanied by increases (decreases) in unemployment-rate inflows and outflows. And at most turning points, movements in the outflow rate clearly lag behind movements in the inflow rate. Three different explanations for these patterns suggest themselves: (1) The pace of labor allocation rises and falls with the unemployment rate, and the lag between the inflow and outflow rate movements reflects the time-consuming nature of finding new jobs. (2) Separations rise and fall with the unemployment rate, but large numbers of job-losers and job-leavers pass through the unemployment pool and then leave the labor force for an extended period of time. (3) Temporary

³An almost identical figure appears in Darby, Haltiwanger, and Plant (1986).

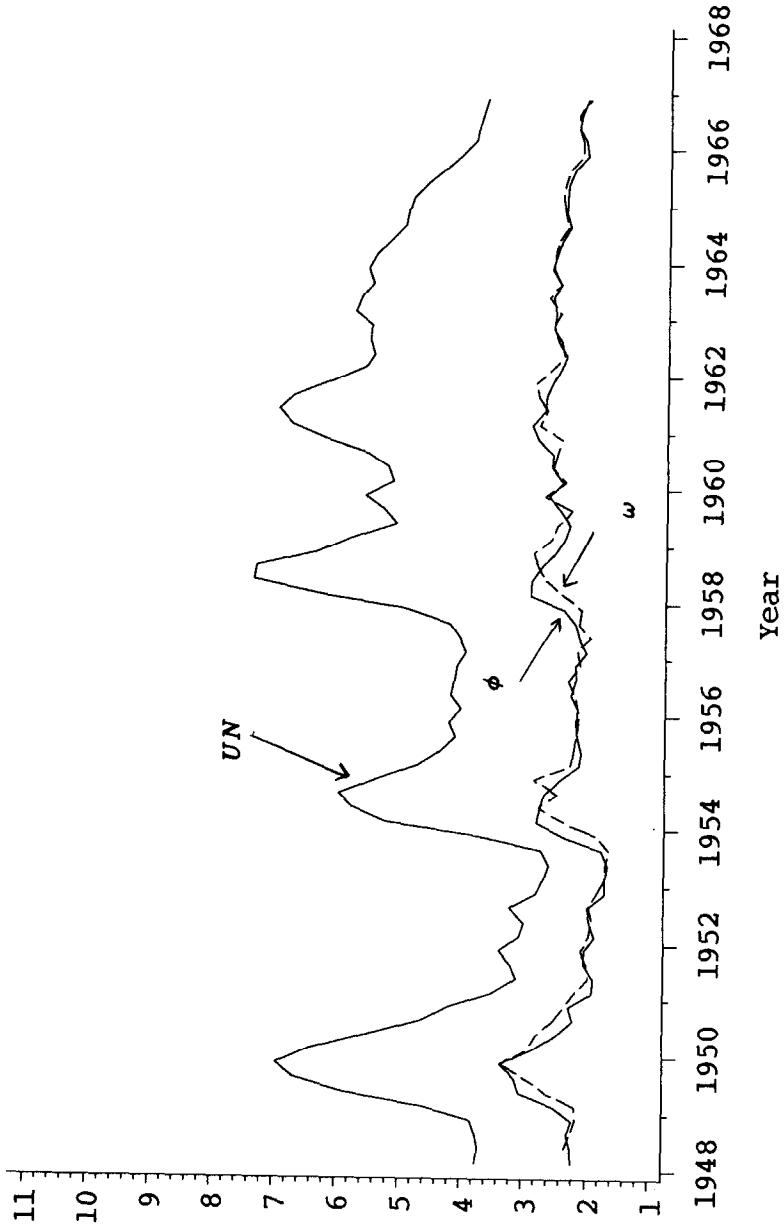
separations rise and fall with the unemployment rate, and the lag between the inflow and outflow rate movements reflects the duration of temporary separations. Explanation (1) implies that the rate of attachments to new jobs rises, even as the unemployment rate rises. This explanation is consistent with the sectoral shifts and reallocation-timing hypotheses. Explanations (2) and (3), though very different, are consistent with the normal business-cycle hypothesis.

Some light is shed on explanation (2) by considering the behavior of the labor-force participation rate. Figure 2 plots the labor-force participation rate, overlaid by a plot of the unemployment outflow rate. This figure provides little support for explanation (2). Indeed, the sample correlation between the labor-force participation rate and the unemployment outflow rate is positive and large--.80. Much of this positive correlation stems from the common upward trend in the two variables since 1970; but, even for the 1948:2 to 1969:4 subsample, the correlation between the labor-force participation rate and the unemployment outflow rate is .11. Over the 1948:2 to 1986:2 (1948:2 to 1969:4) sample, the correlation between the outflow rate and the percent change in the labor-force participation rate is .03 (-.03). While casting some doubt on explanation (2), these results are not conclusive. Suppose, for example, that following an initial wave of separations, out-of-work persons rotate back and forth from unemployed to not in the labor force for an extended period of time, but the labor-force participation rate remains constant. There is no evidence here against this variant of explanation (2).

Clark and Summers (1979), Lilien (1980), Bednarzik (1983), Lilien and Hall (1986), and others use CPS data and manufacturing labor turnover data to assess the contribution of temporary layoff unemployment to total unemployment and to unemployment fluctuations. Temporary layoff unemployment accounts for a small fraction of total unemployment, about one-seventh on average, and a larger fraction of unemployment fluctuations. The literature on temporary layoff unemployment suggests that explanation (3) probably accounts, in part, for the patterns that appear in figure 1. The smallness of temporary layoff unemployment relative to total unemployment suggests, however, that (3) cannot fully explain the patterns in Figure 1.

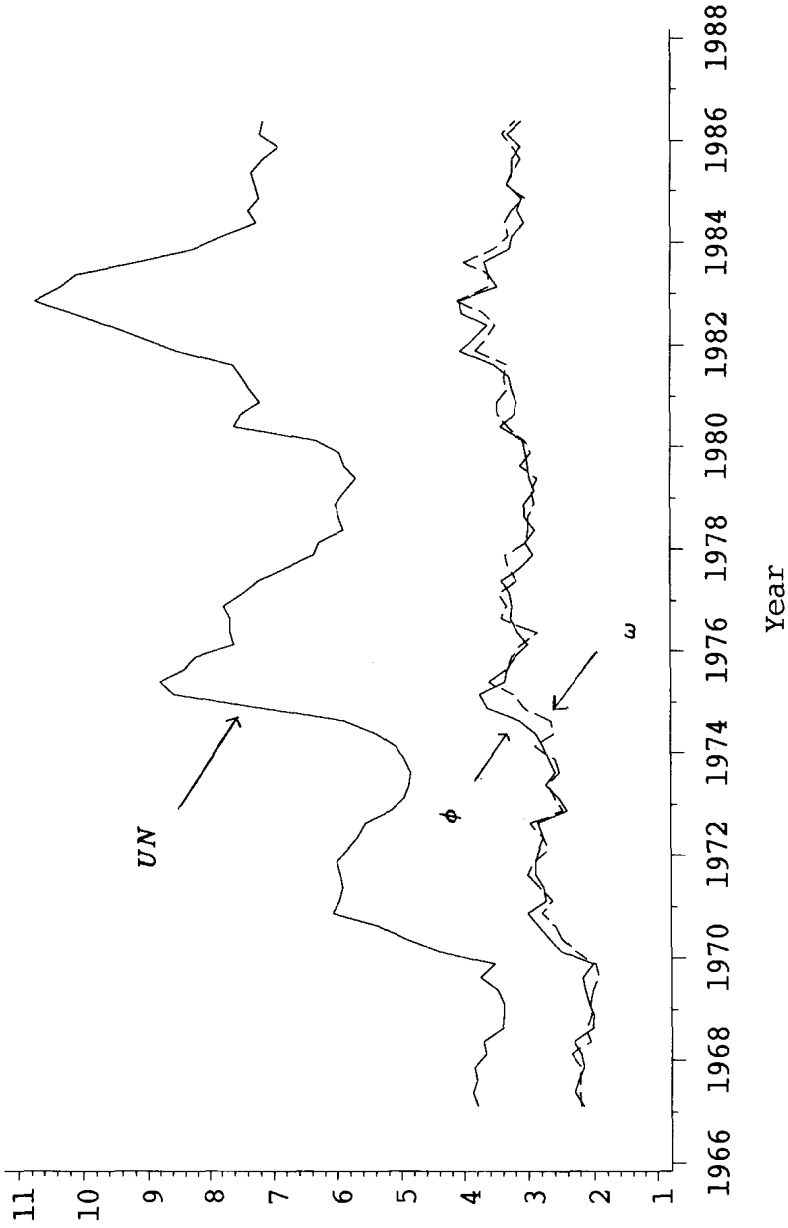
Since 1967, the CPS has identified the "reason for unemployment" by classifying unemployed individuals as new entrants, reentrants, job leavers, job losers on layoff, or other job losers. (In CPS terminology, job losers on layoff encompass "temporary layoff" persons--those who have a

Figure 1
Unemployment Inflow and Outflow Rates, 1948:1 to 1986:2



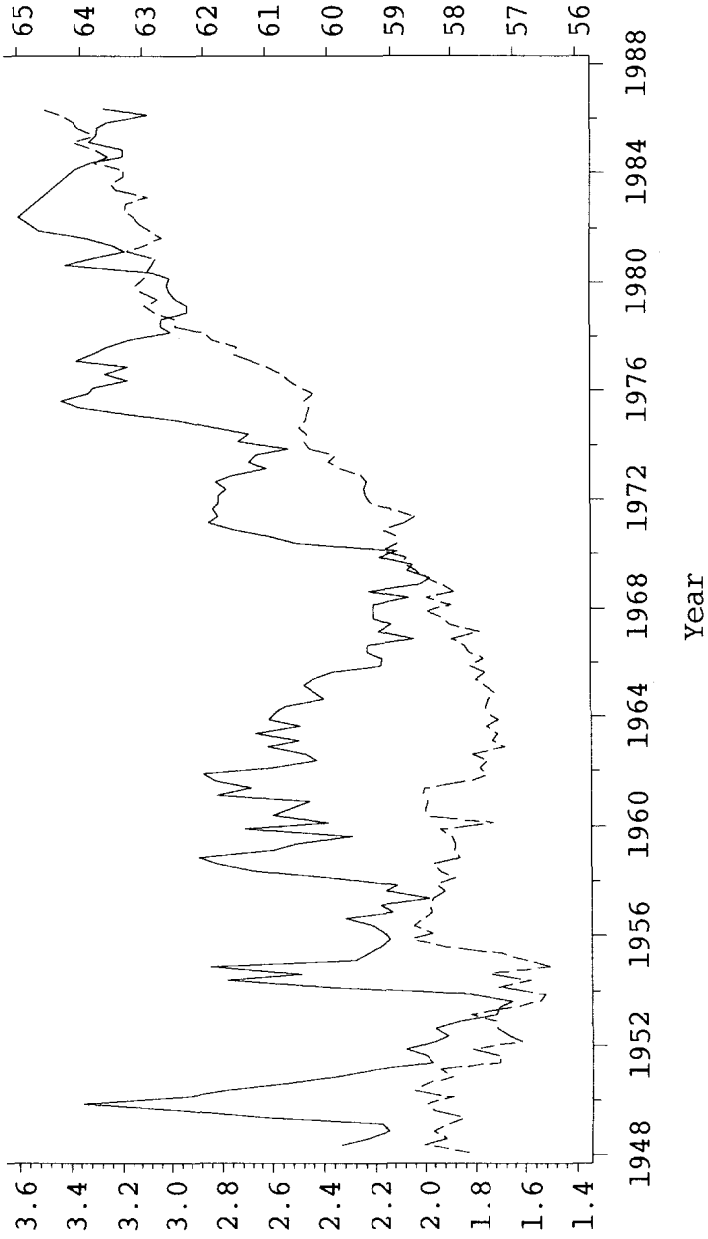
Quarterly averages of monthly, seasonally adjusted data.

Figure 1 (continued)
 Unemployment Inflow and Outflow Rates, 1948:1 to 1986:2



Quarterly averages of monthly, seasonally adjusted data.

Figure 2
Labor Force Participation and Unemployment Outflow Rates, 1948:2-1986:2



Dashed line: labor force participation rate.
 Solid Line: unemployment outflow rate.

definite recall date within thirty days--and persons on indefinite layoff.) Changes in the combined job losers category dominate changes in total unemployment. The distinction between job losers on layoff and other job losers can be used to assess the role of temporary and permanent separations in unemployment fluctuations. Using the CPS data on unemployment by reason, Bednarzik (1983, p. 7) tabulates the rise in the two job losers categories as a percent of the rise in unemployment for post-1967 business-cycle peaks to troughs:

Increase in Job Losers as a Percent of Increase in Unemployed Persons

Time Interval	Total Job Losers	Laid Off Persons	Other Job Losers
December 1969 to November 1970	60.0	22.9	37.1
November 1972 to March 1975	72.6	35.3	37.3
January 1980 to July 1980	82.3	46.3	36.0
July 1981 to November 1982	84.5	31.4	53.1

These figures suggest that changes in both temporary and permanent separations contribute to short-run unemployment-rate increases. Interestingly, the figures indicate that permanent separations played an especially important role in the severe 1981/1982 recession.

Several pieces of evidence indicate that these figures substantially understate the contribution of permanent separations to short-run unemployment-rate fluctuations. First, a large fraction of persons who report themselves as laid-off do not return to their previous job. Using month-to-month matched samples of CPS respondents in 1976, Clark and Summers (1979, p. 49) estimate that only 51% of persons on temporary layoff return to jobs in the same two-digit level industry and occupation. Similarly, based on a month-to-month matched CPS sample from 1976, Bednarzik (1983, pp. 8-9) finds that 37% of laid-off persons in month one who were employed in month two report a change in industry attachment at the three-digit level. Undoubtedly, these figures partly reflect respondent reporting errors and interviewer coding errors, but only 14% of persons employed in consecutive months report changes in industry attachment at the three-digit level. Second, many persons on temporary layoff engage in extensive job-search activity. Based on the May 1976 CPS,

Clark and Summers (p. 48) report that the mean hours of search per month for persons on temporary layoff is 18.3, as compared to 24.9 for all unemployed persons. Using data from a special Methods Development Survey conducted over the April 1981 to December 1982 period, Bednarzik (p. 8) reports that 58% of all persons who report themselves as laid off also report that they looked for work during the preceding four weeks. Third, many CPS respondents interpret the term "layoff" as synonymous with job termination. Bednarzik discusses a special CPS reinterview survey of unemployed persons in July 1982 conducted two weeks after the initial survey. When asked directly, "Do you eventually expect to be called back to the job from which you were on layoff?", almost one-fourth of those persons still unemployed responded in the negative. Combined with the figures in the preceding table, these pieces of evidence indicate that permanent separations constitute the most important component of short-run unemployment-rate fluctuations.

In light of these observations on the relative importance of temporary and permanent separations, and given the absence of support for (2), I conclude that explanation (1) is largely responsible for the patterns in Figure 1. Taken together, these bits of evidence indicate that fluctuations in the pace of labor reallocation are a key aspect of short-run unemployment-rate fluctuations.

B. THE PACE OF LABOR REALLOCATION AND VACANCY MEASURES

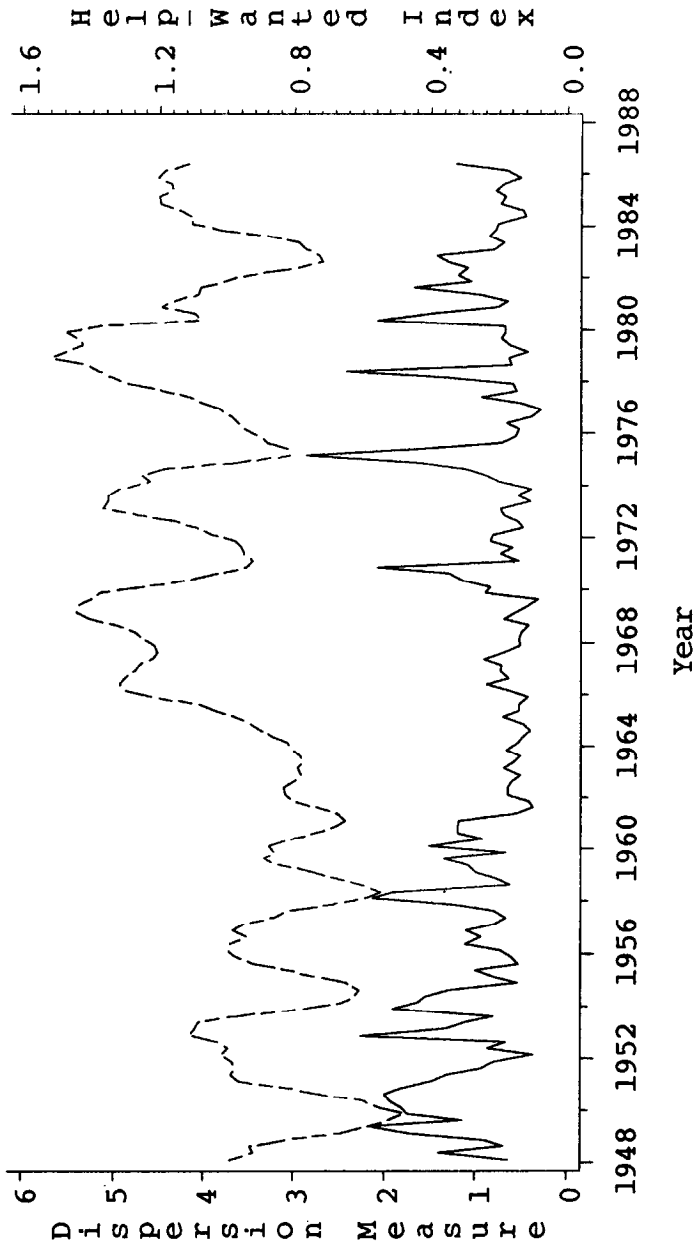
Abraham and Katz (1986) document the negative time-series relationship between vacancy-rate proxies and Lilien-type dispersion measures. They argue that this negative relationship constitutes grounds for rejecting the sectoral shifts hypothesis.

Quarterly data for the dispersion measure and for the normalized help-wanted index are plotted in Figure 3. As in Abraham and Katz, the normalized help-wanted index is computed as the Conference Board's help-wanted index divided by civilian employment. The construction of the dispersion measure is detailed in Section III. Figure 3 depicts the negative relationship between the vacancy-rate proxy and the dispersion measure; the contemporaneous correlation between these two measures is $-.378$.

What can we conclude about the sectoral shifts, reallocation-timing, and normal business-cycle hypotheses from the negative correlation between the vacancy proxy and the dispersion measure? Very little, I argue.

Figure 4 plots the normalized help-wanted index against the

Figure 3
Dispersion Measure and Normalized Help-Wanted Index, 1948:1 to 1986:2



Dashed line: normalized help-wanted index.
 Solid Line: cross-sectoral dispersion in employment growth rates.

unemployment outflow and inflow rates. This figure indicates the difficulty of using the short-run behavior of the vacancy proxy to draw inferences about short-run movements in the pace of labor reallocation and the rate of attachments to new jobs. Periods of high and rising inflow and outflow rates coincide with periods of a declining help-wanted index, and vice versa. Upon reflection, this finding seems unsurprising. Many jobs are filled without any vacancy spell (and many establishments seem always to have vacancies), but we have little direct evidence about cyclic variation in the extent to which jobs are filled without a vacancy spell. More generally, we have little direct evidence about the cyclic variation of vacancy durations (including jobs filled with zero vacancy duration). The indirect evidence provided by Figure 4 suggests the following stylized account: During periods with a low stock of existing vacancies, there is a rapid reallocation of labor across jobs. Also, during periods with a low vacancy stock, the flow rate of new job vacancies (again, including jobs filled with a zero vacancy duration) is unusually high. The sharp decline in the mean vacancy duration more than offsets the effect of an increased flow of new vacancies on the existing stock.

The key point of this stylized account of vacancy fluctuations is that the cyclic behavior of vacancy flows can differ dramatically from the cyclic behavior of the vacancy stock. To fully appreciate this point, consider a direct analogy to the cyclic behavior of quits. Time-series data are available on both the quit rate, the rate at which workers initiate separations from employers, and the stock of quit unemployment. The quit rate is the analog to the flow of new vacancies, and the stock of quit unemployment is the analog to the vacancy stock. It is well-known that the quit rate is highly procyclic; see, for example, Table 3 in Prescott et al. (1983). However, one cannot hope to discern the procyclicality of the quit rate from time-series data on the stock of quit unemployment, as the following table shows.

Number of Unemployed Persons Who
Report Themselves as Job Leavers (in thousands)

Year	1967	1968	1969	1970	1971	1972	1973	1974	1975	1976
Number	438	431	436	550	590	641	683	768	827	903
Year	1977	1978	1979	1980	1981	1982	1983	1984	1985	
Number	909	874	880	891	923	840	830	823	877	

Evidently, the mean unemployment spell duration associated with quits (including quits that involve no unemployment) moves inversely to the quit rate, resulting in roughly acyclic movements in the stock of quit unemployment. Hence, we find a dramatic difference between the cyclic behavior of the quit stock measure and the cyclic behavior of the quit flow measure.

In conclusion, the prediction of the sectoral shifts hypothesis (and the reallocation-timing hypothesis) that a rapid rate of attachments to new jobs accompanies cyclically high unemployment rates is fully consistent with the observed behavior of vacancy stock measures.

III. CONDITIONING AND STAGE-OF-BUSINESS-CYCLE EFFECTS IN THE RELATIONSHIP BETWEEN UNEMPLOYMENT AND LABOR REALLOCATION

In this section of the paper, I adopt the sectoral shifts interpretation of Lilien's empirical finding and consider the following question: What further implications does the sectoral shifts hypothesis carry for the behavior of unemployment? I develop and test two implications that have not been noted in previous research. Elsewhere (Davis, 1986a,b) I develop and test the implication that allocative disturbances cause temporally asymmetric unemployment rate fluctuations.

A. CONDITIONING ON PAST PATTERNS AND MAGNITUDES OF LABOR REALLOCATION

In motivating the sectoral shifts hypothesis, Lilien (1982, 1983) and Lilien and Hall (1986) emphasize that long-term attachments to specific sectors can impede or otherwise influence the process of labor reallocation. Long-term attachments to specific sectors stem from sector-specific human capital and job contacts that accumulate and depreciate slowly, lump-sum sector-switching costs, and from sectoral wage (or utility) differentials that reflect noncompetitive forces (see Hall (1975)). These short-run barriers to labor mobility imply that the contemporaneous relationship between intersectoral labor reallocation and aggregate unemployment depends, in part, on past patterns and magnitudes of labor reallocation. A laid-off worker who attaches some probability to an improvement in his sector's fortunes might not immediately seek to switch sectors. Such a worker is available for immediate recall in the event that his sector's fortunes improve. Hamilton's (1986) model exhibits this type

Figure 4
Normalized Help-Wanted Index and Unemployment Flow Rates, 1948:1 to 1986:2

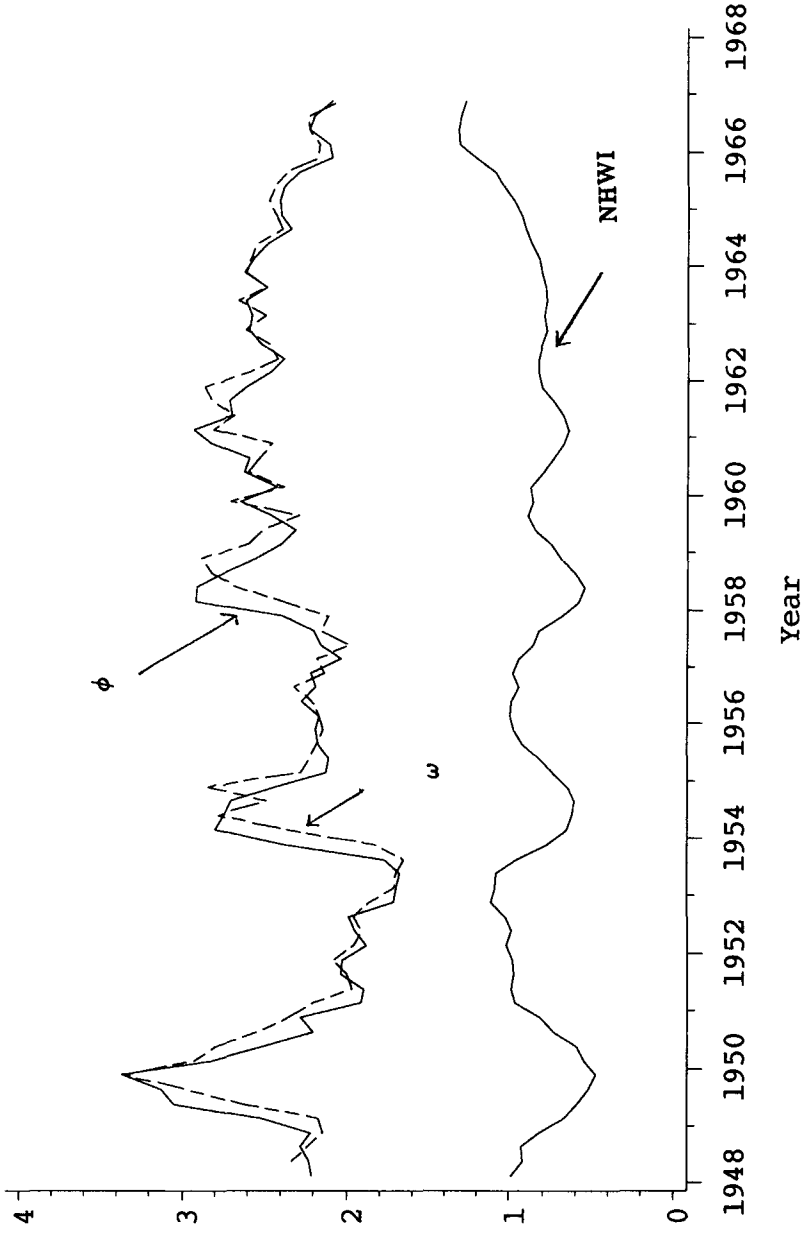
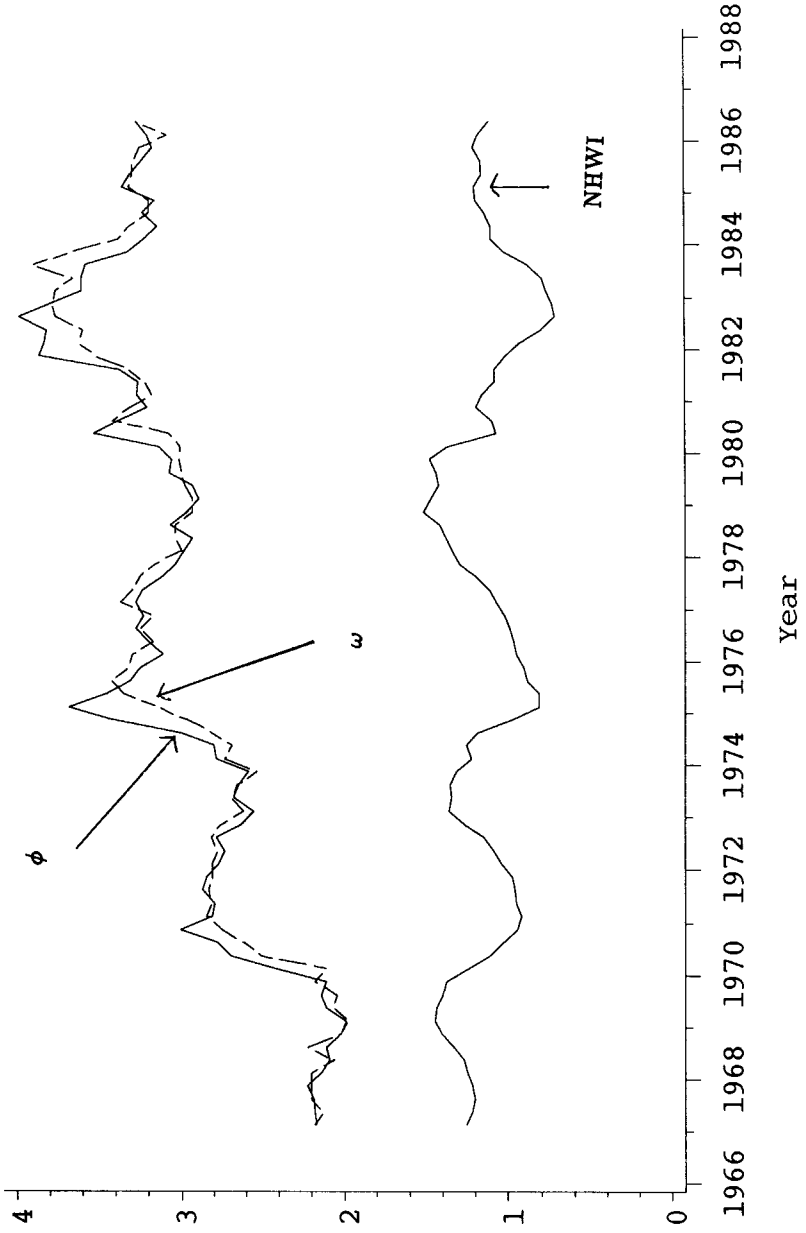


Figure 4 (continued)
Normalized Help-Wanted Index and Unemployment Flow Rates, 1948:1 to 1986:2



of unemployment. A worker who has switched sectors will find it relatively easy to return to his former sector, if its fortunes improve, because of his accumulated sector-specific capital, job contacts, seniority, etc.

The potential importance of conditioning on past patterns of labor reallocation is easily seen by considering a concrete example. The relative price of crude petroleum fell 49.9% over the three-month span ending in May 1986 (calculated from the producer price index for crude petroleum, deflated by the all-commodities index). This tremendous relative price decline benefits many geographic, industrial, and occupational sectors of the United States economy while adversely affecting other sectors. Recent unemployment-rate increases in Louisiana, Oklahoma, and Texas--states whose economic fortunes are closely tied to the economic fortunes of the petroleum industry--reflect the adjustment process by which labor reallocates away from petroleum-dependent industries, regions, and occupations and toward other sectors. Recognizing this recent pattern of labor reallocation is clearly relevant to predicting the unemployment response to any further oil price shock or other allocative disturbance. In particular, an allocative disturbance in the near future that favorably affects petroleum-dependent sectors would prompt a reallocation of labor towards these sectors accompanied by little increase, or possibly a decrease, in unemployment. In contrast, an allocative disturbance in the near future that further adversely affects these sectors would significantly increase unemployment. In terms of the conceptual framework introduced in Section I, a favorable allocative disturbance is a fortuitous event that shifts the target allocation of labor resources in the direction of the actual allocation.

This example, and others like it, suggests that empirical unemployment-rate models could be improved by conditioning on past patterns of labor reallocation. Implementing this conditioning strategy clearly requires some quantification of the concept, "past pattern of labor reallocation." I introduce a crude, but workable, first approach to quantifying this concept.

Existing empirical studies in the sectoral-shifts literature use the following dispersion measure (or some variant) to measure the magnitude of (net) intersectoral labor reallocation:

$$\sigma_t = \left(\sum_{i=1}^N \left(\frac{x_{it}}{\bar{x}_t} \right) (\Delta_1 x_{it} - \Delta_1 \bar{x}_t)^2 \right)^{\frac{1}{2}} \quad (1)$$

where

x_{it} = employment in sector i at time t ,

X_t = aggregate employment at time t ,

$\Delta_j x_{it} = \ln x_{it} - \ln x_{i,t-j}$, and

N = the number of labor-market sectors, broken down by industrial classification or geographic region.⁴

Recognizing σ_t as the (square root of the) weighted cross-sectoral variance in employment growth rates suggests the use of weighted cross-sectoral covariance measures as simple devices for capturing the effects of past patterns of labor reallocation. I use the following type of cross-sectoral covariance measure:

$$\sigma_{t,j}^H = \sum_{i=1}^N \left(\frac{x_{it}}{X_t} \right) (\Delta_1 x_{it} - \Delta_1 X_t) (\Delta_j x_{i,t-1} - \Delta_j X_{t-1}), \quad j = 1, 2, \dots, J. \quad (2)$$

$\sigma_{t,j}^H$ indexes the time t direction of labor reallocation over a one-period horizon relative to the $t-1$ direction over a j -period horizon. This index provides a workable method of conditioning on past patterns of labor reallocation in time-series data. Relatively large (small) values for $\sigma_{t,1}^H, \sigma_{t,2}^H, \dots$ indicate that the time t direction of labor reallocation reinforces (reverses) past patterns of labor reallocation.⁵ Reinforcement (reversal) of recent past patterns of labor reallocation exacerbates (mitigates) skill, location, and informational mismatches between workers and firms. Thus, the sectoral-shifts hypothesis predicts a positive partial correlation between the directional indexes and the economywide unemployment rate.

Past magnitudes of labor reallocation will be reflected in the relationship of unemployment to lagged values of the dispersion measure, σ_t . Workers who suffer large wage declines as a result of allocative disturbances that cause them to switch sectors may exhibit a much weaker attachment to employment than they did in their initial sector. Their new market wage will be closer to their (new) reservation wage. Also, workers

⁴In the quarterly data, the dispersion measure is based on an eleven-industry decomposition of nonagricultural employment: state and local government, federal government, wholesale trade, retail trade, the FIRE group, durable-goods manufacturing, nondurable-goods manufacturing, transportation and public utilities, contract construction, mining, and services.

⁵Long-term, ongoing changes in the sectoral distribution of the labor force raise the average value of the cross-sectoral covariance measures.

induced to switch sectors by allocative disturbances may repeat the matching and turnover process (accompanied by frequent unemployment spells) that is typical of new entrants into the labor-market. Akerlof and Main (1981) and Hall (1982) use job tenure data from special Current Population Surveys to document the importance of near-lifetime jobs--jobs that last at least twenty years--in the United States economy. Despite the importance of long-tenure jobs, the probability that a given new job will become a near-lifetime job is less than six percent for all age groups (Hall, 1982, p. 720). These observations suggest that allocative disturbances, by disrupting long-term worker-job matches, can increase the unemployment propensities of affected workers for several years.

The hypothesis suggested here--that past patterns and magnitudes of labor reallocation influence current unemployment behavior--in some respects has the same flavor as the "persistence hypothesis" of labor force participation advanced by Clark and Summers (1982). In characterizing the persistence hypothesis, Clark and Summers (p. 826) write:

In this view, past work experience is a key determinant of current employment status. Because of high separation costs and costs of finding new employment, those who are employed tend to remain employed. Persistence of employment might also be rationalized on human capital grounds. Those who are employed longer tend to accumulate more human capital, which raises the return to work in the future relative to leisure. Those out of the labour force may also develop household-specific capital or commitments (i.e., children) which reduce the return to working relative to remaining outside the labour force.

Clark and Summers present evidence that persistence effects are an important aspect of movements in the labor-force participation rate.

I first test for evidence of conditioning effects in the relationship between unemployment and labor reallocation using post-World War II quarterly data. Below, I report some results based on annual data over a longer sample period.

TESTS BASED ON QUARTERLY DATA

Table 1 reports the results of estimating a benchmark specification of an unemployment-rate equation that is similar to specifications estimated in my 1986(b) paper:

Table 1

Joint Nonlinear Least Squares Estimation of the Unemployment Equation

Sample: 1953:2 - 1986:2

$$UN_t = c + \tau DUM74_t + \sum_{i=0}^{12} \delta_i DMA_{t-i} + \sum_{i=0}^9 \beta_i DMU_{t-i} + \sum_{i=0}^{12} \gamma_i \sigma_{t-i} + \rho_1 u_{t-1} + \rho_2 u_{t-2} + \epsilon_t$$

Parameter	Estimate	Standard Error	Marginal Significance Level
c	.842	(.641)	.192
τ	.688	(.258)	.009
δ_0	-.056	(.071)	.427
δ_1	-.194	(.096)	.046
δ_2	-.185	(.104)	.080
δ_3	-.288	(.106)	.008
δ_4	-.322	(.106)	.003
δ_5	-.352	(.104)	.001
δ_6	-.353	(.109)	.002
δ_7	-.382	(.107)	.001
δ_8	-.372	(.107)	.001
δ_9	-.332	(.106)	.002
δ_{10}	-.278	(.101)	.007
δ_{11}	-.180	(.088)	.045
δ_{12}	-.071	(.062)	.255
β_0	-.107	(.050)	.036
β_1	-.147	(.082)	.076
β_2	-.284	(.101)	.006
β_3	-.312	(.109)	.005
β_4	-.315	(.113)	.007
β_5	-.334	(.115)	.005
β_6	-.258	(.112)	.024
β_7	-.267	(.105)	.013
β_8	-.187	(.091)	.043

Table 1 continued

Parameter	Estimate	Standard Error	Marginal Significance Level
β_9	-.029	(.062)	.640
γ_0	.361	(.063)	.000
γ_1	.559	(.084)	.000
γ_2	.384	(.092)	.000
γ_3	.335	(.096)	.001
γ_4	.138	(.096)	.154
γ_5	.113	(.096)	.242
γ_6	.224	(.096)	.021
γ_7	.288	(.094)	.003
γ_8	.218	(.096)	.027
γ_9	.203	(.095)	.036
γ_{10}	.138	(.095)	.148
γ_{11}	.151	(.086)	.181
γ_{12}	.252	(.065)	.000
ρ_1	1.356	(.090)	.000
ρ_2	-.554	(.092)	.000
$R^2 = .9870$		S.E.E. = .2436	

- Notes: i. The equation is estimated jointly with (4) by nonlinear least squares.
- ii. S.E.E. is the square root of the mean squared residual.
- iii. The estimated standard errors are based on a degree of freedom correction that attributes one-half degree of freedom to each equation for the shared parameters in the model; i.e., the β 's, and the δ 's.
- iv. Marginal significance level is based on a two-tailed asymptotic t-test of the null hypothesis that the parameter of interest is zero.
- v. The dispersion measure and DM are scaled up by a factor of 10 and 100, respectively, prior to estimation.

$$UN_t = c + \tau DUM74_t + \sum_{i=0}^{12} \delta_i DMA_{t-i} + \sum_{i=0}^9 \beta_i DMU_{t-i} + \sum_{i=0}^{12} \gamma_i \sigma_{t-i} + \rho_1 u_{t-1} + \rho_2 u_{t-2} + \varepsilon_t, \quad (3)$$

where UN is the quarterly, seasonally adjusted, civilian unemployment rate; DUM74 is a dummy variable that equals zero prior to 1974 and one thereafter; DMU is the unanticipated component of the money-supply growth rate, based on one-quarter-ahead forecasts; DMA is the anticipated (net of estimated trend) money-supply growth rate; σ , the dispersion measure described by (1), is the proxy for the effects of allocative disturbances; and the error term obeys a second-order autoregressive process. Detailed descriptions of the data and their sources appear in the Appendix.

The forecasting equation for the money supply growth rate,

$$DM_t = a + b TIME_t + \sum_{i=1}^{12} n_i DM_{t-i} + \sum_{i=1}^4 m_i TBILL_{t-i} + \sum_{i=1}^4 l_i UN_{t-i} + v_t, \quad (4)$$

contains twelve lags of the money-supply growth rate, measured as the change in the log of the seasonally adjusted M1 stock, four lags of the 90-day Treasury bill rate, four lags of the civilian unemployment rate, a constant, and a linear time trend. $DMU_t = DM_t - \hat{DM}_t$ and $DMA_t = \hat{DM}_t - b TIME_t$, where \hat{DM}_t is the predicted value of the money-supply growth rate in (4).

Equations (3) and (4) are estimated jointly by nonlinear least squares, imposing the cross-equation restrictions implied by the forecasting mechanism, and permitting the variance of the disturbance terms to differ across equations. The β 's and γ 's in (3) are identified under the assumptions that $cov(\varepsilon_t, v_{t-j}) = 0$ for $j \geq 0$ and that lagged values of the Treasury bill rate do not enter the unemployment-rate equation except through the anticipated money-supply growth rate terms.

The estimation results in Table 1 for the 1953:2 to 1986:2 sample period are similar to results, based on a shorter sample period, reported in my 1986(b) paper.⁶ With the exception of an unexplained jump of seven-

⁶In my earlier paper, I have estimated variants of (3) and (4) with a different unemployment-rate measure, different dispersion measures, a different specification of the forecasting equation, and different specifications of the unemployment-rate equation. The numerous changes I considered do not affect the pattern of results exhibited in Table 1.

tenths of a percentage point around 1974, the regression model accounts for the long-term as well as short-term movements in the unemployment rate. When a linear time-trend variable (incrementing by 1 each quarter) is added to (3), it has a coefficient estimate of .004 and a standard error estimate of .015. When a linear trend that begins in 1974 is added to the model, it has a coefficient estimate of -.020 and a standard error estimate of .019. Neither the parameter estimates nor the estimated standard errors in Table 1 are appreciably affected by the inclusion of the time trend variables.

Both types of monetary disturbances and the dispersion measure show statistically significant effects on the unemployment rate. Anticipated and unanticipated monetary disturbances are negatively related to unemployment rate movements. The distributed lags on the monetary disturbance measures are hump-shaped and indicate a peak effect of DMU on the unemployment rate at a lag of three to five quarters. The peak effect of DMA on the unemployment rate extends over several additional quarters. The dispersion measure is positively correlated with the unemployment rate as predicted by the three competing hypotheses. A likelihood ratio test, performed using Gallant and Jorgenson's (1979) T^0 statistic, of the joint hypothesis that $\gamma_0 = \gamma_1 = \dots = \gamma_{12} = 0$ is overwhelmingly rejected. Under the null hypothesis, the test statistic is asymptotically Chi-square distributed with thirteen degrees of freedom. The critical value for rejection of the null at the 1% level is 27.7, and the computed test statistic value is 79.7.

The strong effects of lagged values of the dispersion measure lend support to the notion that allocative disturbances have effects on workers' unemployment propensities that persist well beyond an initial spell of unemployment. Thus, according to this interpretation of Table 1, unusually large magnitudes of labor reallocation in the past imply higher current

unemployment rates.⁷

Figure 5 plots the movements in the unemployment rate explained by movements in the dispersion measure, overlaid by a plot of the actual unemployment-rate series. The simulated series is generated by letting σ and its lags assume their sample values, while fixing all other regressors in (3) at their sample means. Using these values for the regressors, the parameter estimates in Table 1 imply the series plotted in Figure 5. This figure shows that movements in the dispersion measure account for a large portion of movements in the unemployment rate, controlling for the effects of anticipated and unanticipated monetary disturbances.

Table 3 reports the results of modifying the benchmark specification (3) by the inclusion of the cross-sectoral covariance measure, $\sigma_{t,j}^H$, for various values of j . Not surprisingly, the observations on the covariance measures are highly collinear for different values of j . See Table 2, which reports summary statistics and correlations for the variance and covariance measures for selected values of j and at selected lags.

The basic messages of Table 3 can be summarized as follows. First, there is clear evidence of a statistically significant, positive relationship between the unemployment rate and the covariance measures, as

⁷The coefficients on the dispersion measure and its lags exhibit a noticeably wavy pattern in Table 1. This wavy pattern is not peculiar to the specification estimated in Table 1; it emerges in all specifications that I have estimated with quarterly data. This wavy pattern appears not to stem from the sources of spurious correlation between σ and the unemployment rate identified by Abraham and Katz. I have reestimated variants of (3) using dispersion measures that are "purged" of the influence of aggregate-demand disturbances, and the wavy pattern reemerges quite strongly. See Table 4 in my 1986(b) paper. This wavy pattern may reflect the dynamic response pattern of unemployment to unobserved exogenous impulses. To see this point, consider the following general linear representation of a structural model that links σ impulses to unemployment:

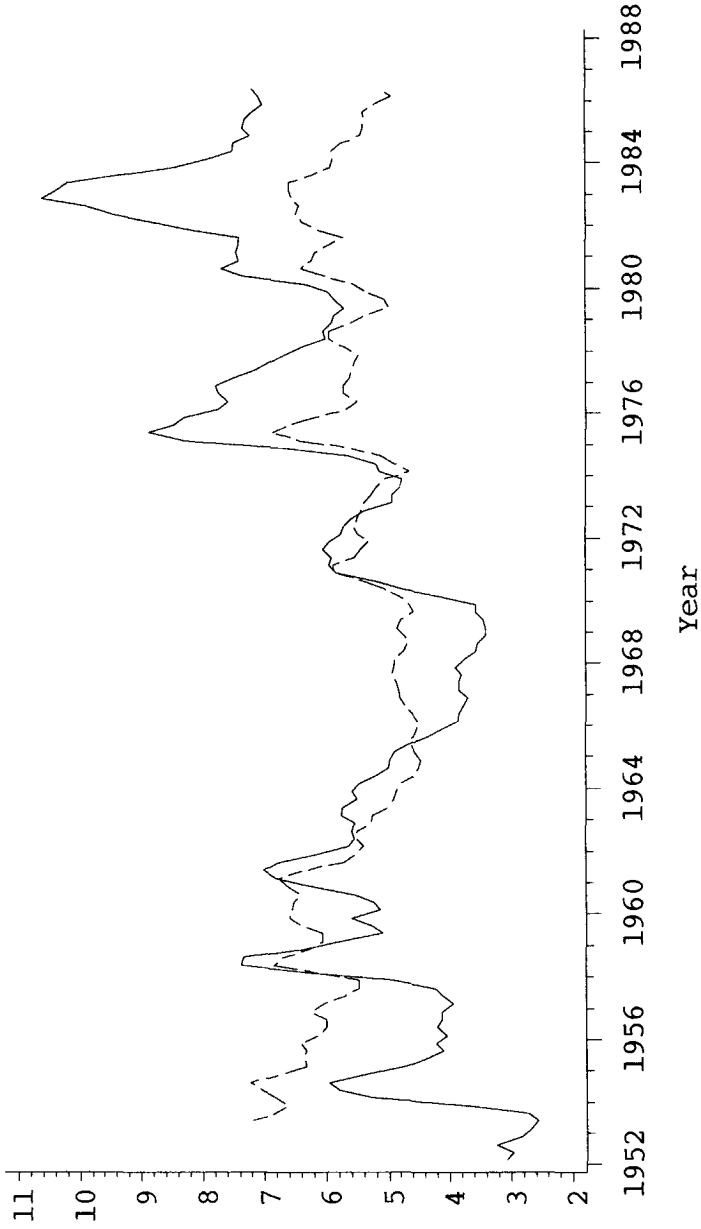
$$A(L)UNV_t = B(L)\sigma_t + C(L)^{-1}\epsilon_t.$$

Here, $B(L)$ describes the dynamic response of unemployment to the σ impulses, $C(L)^{-1}$ describes the dynamic response of unemployment to unobserved exogenous impulses, and $A(L)$ represents endogenous sources of dynamic response to any unemployment impulse. Since sector-specific forms of human capital depreciate over time, the sectoral shifts hypothesis suggests that the coefficients in the lag operator $B(L)$ decline monotonically in lag length. Unfortunately, $B(L)$ is not identified in the estimated equation. To construct the least-squares estimation criterion function, one uses the error term:

$$\epsilon_t = C(L)D(L)UN_t - C(L)B(L)\sigma_t.$$

From this representation, it is clear that the estimated coefficients on the dispersion measure and its lags reflect the dynamic unemployment response to σ and to the unobserved exogenous impulses.

Figure 5
Actual and Simulated Unemployment Rate Series, 1953:2 to 1986:2



The solid line represents the actual unemployment rate.
The dashed line represents the simulated unemployment rate.

Table 2

Summary Statistics and Correlations for Selected Cross-Sectoral Variance and Covariance Measures
1953:2 to 1986:2

	σ_t	σ_{t-1}	σ_{t-3}	σ_{t-6}	σ_{t-9}	σ_{t-12}	$\sigma_{t,1}^H$	$\sigma_{t,3}^H$	$\sigma_{t,6}^H$	$\sigma_{t,9}^H$	$\sigma_{t,12}^H$
σ_t	.8570 (.4507)										
σ_{t-1}		0.489 .8578 (.4515)									
σ_{t-3}			.8709 (.4663)								
σ_{t-6}				.8700 (.4679)							
σ_{t-9}					.8854 (.4694)						
σ_{t-12}						.9107 (.4940)					
$\sigma_{t,1}^H$							0.511				
$\sigma_{t,3}^H$							0.496	0.520			
$\sigma_{t,6}^H$							0.089	0.504	0.483		
$\sigma_{t,9}^H$							0.089	-0.004	0.435	0.420	
$\sigma_{t,12}^H$							0.030	0.004	-0.131	-0.115	0.285
							-0.051	-0.082	-0.043	0.043	0.377
							-0.022	-0.062	-0.145	-0.116	-0.112
							0.044 (0.084)	0.894	0.673	0.585	0.064
							0.087 (0.168)	0.087 (0.168)	0.836	0.720	0.635
									0.102 (0.201)	0.874	0.784
										0.126 (0.221)	0.896
											0.155 (0.273)

Notes:

- i. Off-diagonal entries show the Pearson correlation between the row and column variables.
- ii. Sample means appear on the diagonal with corresponding standard deviations in parentheses.

Table 3

Estimated Effects of the Cross-Sectoral Covariance Measures
Sample: 1953:2- 1986:2

Benchmark Specification:

$$U_t = c + \tau \text{DUM74}_t + \sum_{i=0}^{12} \delta_i \text{DMA}_{t-i} + \sum_{i=0}^9 \beta_i \text{DMU}_{t-i} \\ + \sum_{i=0}^{12} \gamma_i \sigma_{t-i} + \rho_1 u_{t-1} + \rho_2 u_{t-2} + \epsilon_t$$

Variable(s) Added to the Benchmark Specification	Coefficient Estimate	Estimated Standard Error	Marginal Significance Level
σ_1^H	.598	.222	.009
σ_6^H	.279	.134	.040
σ_{12}^H	.166	.078	.038
σ_1^H	.483	.251	.058
σ_{12}^H	.089	.088	.317
σ_1^H	.519	.291	.078
σ_6^H	-.062	.215	.774
σ_{10}^H	.136	.132	.308
σ_1^H	.639	.364	.083
σ_4^H	-.108	.229	.639
σ_8^H	-.129	.225	.568
σ_{12}^H	.172	.146	.245

Note: (See, also, notes to Table 1).

- i. Coefficient and standard error estimates on the variables in the benchmark specification are not appreciably affected by inclusion of the covariance measures.

predicted. For example, reading across the first panel of the table, when $\sigma_{t,1}^H$ is added to the benchmark unemployment-rate specification, its coefficient estimate is .60 with an estimated standard error of .22. Second, due to the collinearity problem, it is difficult to assess whether the statistical significance of the covariance measures merely reflects a pattern of short-term layoff and recall or a pattern of workers returning to an initial sector after an extended absence. The results in Table 3 favor the former view. Third, the magnitude of the effect of the covariance measures on the unemployment rate is quite small. Again using the first panel of Table 3 as an example, the coefficient estimate and the range of sample observations on $\sigma_{t,1}^H$ imply a range of movements in the unemployment rate of less than one-half of a percentage point.

B. STAGE-OF-BUSINESS-CYCLE EFFECTS IN THE RELATION BETWEEN UNEMPLOYMENT AND LABOR REALLOCATION

Regression models employed in previous empirical work on the sectoral shifts hypothesis impose a fixed relation, independent of the stage of the business cycle, between unemployment and cross-sectoral dispersion measures. If, however, dispersion measures proxy for the magnitude of labor reallocation, then there are several reasons for potentially important stage-of-business-cycle effects in the relationship between these measures and the unemployment rate.

First, if the mobility technology available to workers permits substitution (at a cost) away from longer unemployment-spell durations, then the relation between unemployment and labor reallocation varies systematically with the opportunity cost of unemployment. This opportunity cost, in turn, varies systematically over the business cycle. Thus, given that a worker moves during good times (bad times)--when the value of foregone output is high (low)--he has an incentive to shorten (lengthen)

the expected duration of his unemployment spell.⁸ Any tendency by the government during bad times to temporarily extend the eligibility period for the receipt of unemployment insurance benefits would reinforce this effect. Note that the substitution into or out of time spent unemployed--given that the worker engages in mobility--is quite distinct from the intertemporal substitution of labor mobility stressed by the reallocation-timing hypothesis. Substitution into or out of time-consuming mobility and unemployment could be important even in a model with an exogenous timing of unemployment inflows.

Second, even if each individual's unemployment exit rate is insensitive to cyclic factors, the composition of unemployment will tend to shift towards low-exit-rate individuals during recessions. Recessions are characterized by sharp increases in unemployment inflow rates--recall Figure 1. To understand the implications of this observation, suppose that individuals are heterogeneous in terms of their unemployment exit probabilities (or exit-probability functions). If, at the onset of a recession, the unemployment entry rate increases equiproportionately for all types of individuals, then the average unemployment exit rate will fall shortly thereafter and remain at an unusually high level for a period of time. Since high-exit-rate individuals leave unemployment quickly, the unemployment rate for these types of individuals will soon return to a lower level commensurate with the normal entry rates for these types. This process takes longer for low-exit-rate individuals--hence, the shift in the composition of the unemployed pool toward low-exit-rate individuals during recessions.⁹ Of course, recessions are probably not well-characterized by

⁸Expected unemployment-spell durations need not, in general, decrease in response to a rightward shift in the wage-offer distribution. See Flinn and Heckman (1983), Burdett and Ondrich (1985), and references therein. These authors consider the optimal search strategy confronting infinitely lived, risk neutral workers facing exogenous offer arrival rates and wage offer distributions in a stationary environment. They describe conditions under which a rightward shift in the wage-offer distribution or an increase in the offer arrival rate decreases the expected duration of unemployment. These theoretical analyses are not directly applicable to the discussion in the text, because I have in mind a situation where the offer arrival rate is a choice variable and the environment is not stationary. My statements in the text about the effect of fluctuations in the opportunity cost of unemployment on expected unemployment duration are only reasonable conjectures--I am not aware of any proof that they are true.

⁹This argument and its implications are developed much more fully in Darby, Haltiwanger, and Plant (1985). The Darby et al. paper and its sequel also contain evidence that heterogeneous exit rates are empirically important and that much of the cyclic variation in exit rates can be explained by cyclic variation in inflow rates.

equiproportionate increases in unemployment entry rates for all types of individuals. If high-exit-rate individuals experience relatively greater unemployment entry rates during recessions, then the cyclic sorting effect described above could be offset or reversed. Note, also, that given heterogeneous exit rates, the cyclic sorting explanation for a stage-of-the-cycle effect in the relationship between dispersion measures and the unemployment rate follows from both the sectoral--shifts and the reallocation--timing hypotheses.

Third, some recent models of unemployment--e.g., Diamond (1984) and Howitt (1985)--emphasize externality feedbacks from aggregate conditions to the labor mobility technology available to individuals. Although the analysis of these models is usually restricted to comparison across alternative steady states, some authors have appealed to these models to argue that the available mobility technology varies systematically over the business cycle, becoming more (less) attractive when aggregate conditions are favorable (unfavorable). If important, these external effects in the labor reallocation process could also generate a stage-of-the-cycle effect.

If the value of production foregone due to labor reallocation is procyclic, then all three of these effects imply the testable hypothesis: a given amount of labor reallocation results in less (more) unemployment when aggregate conditions are relatively favorable (unfavorable). This proposition is tested here by comparing the relationship between the dispersion measure and the unemployment rate at different stages of the business cycle using NBER reference-cycle dating of expansions and recessions.

TESTS BASED ON QUARTERLY DATA

Table 4 reports tests for stage-of-business-cycle effects in quarterly data. These tests are performed by adding various recession interaction variables to the benchmark specification. For example, the first row panel in Table 4 reports the coefficient estimate and marginal significance level on the product of RECESS and $(\gamma_0\sigma_0 + \gamma_1\sigma_{-1})$. RECESS equals the number of months to recession during the quarter divided by three. In line with the results in Table 3, the benchmark specification now includes the cross-sectoral covariance at a one-period horizon.

The interaction variable point estimate in the first row panel of Table 1 is .172, which says that a given amount of labor reallocation is associated with roughly 17% more measured unemployment during recessions than at other times. The marginal significance level on this positive

Table 4

Estimated Stage-of-Business Cycle Effects in the Unemployment Equation

Sample: 1953:2 - 1986:2

Benchmark Specification

$$UN_t = c + \tau DUM74_t + \sum_{i=0}^{12} \delta_i DMA_{t-i} + \sum_{i=0}^9 \beta_i DMU_{t-i} \\ + \sum_{i=0}^{12} \gamma_i \sigma_{t-i} + \theta \sigma_{t,1}^H + \rho_1 u_{t-1} + \rho_2 u_{t-2} + \epsilon_t$$

Variable(s) Added to the Benchmark Specification	Coefficient Estimate	Estimated Standard Error	Marginal Significance Level
RECESS($\sum_{i=0}^1 \gamma_i \sigma_{-i}$)	.172	.088	.054
RECESS($\sum_{i=0}^5 \gamma_i \sigma_{-i}$)	.046	.041	.264
RECESS($\sum_{i=0}^{10} \gamma_i \sigma_{-i}$)	.077	.055	.165
RECESS($\sum_{i=0}^{12} \gamma_i \sigma_{-i}$)	.038	.036	.293
RECESS($\sum_{i=0}^1 \beta_i DMU_{-i}$)	.936	1.122	.406
RECESS($\sum_{i=0}^9 \beta_i DMU_{-i}$)		Parameter estimates did not converge.	
RECESS($\sum_{i=0}^1 \delta_i DMA_{-i}$)		Parameter estimates did not converge.	
RECESS($\sum_{i=0}^{12} \delta_i DMA_{-i}$)	.063	.060	.295

Notes: (See, also, notes to Table 1.)

- i. For those models with convergent parameter estimates, inclusion of the interaction variables does not greatly affect the coefficient and standard error estimates on the benchmark specification variables.

effect is about .05, providing strong evidence against the null hypothesis of no stage-of-cycle effect. However, this result is sensitive to the exact specification of the interaction variable. The second- through fourth-row panels in Table 4 display results for interaction variables that contain sums over successively longer lags on the cross-sectoral dispersion measure. The point estimates on these interaction variables are uniformly positive but substantially smaller than the point estimate in row one. The marginal significance levels indicate very weak evidence against the null hypothesis. In sum, these first four rows of Table 4 provide modest evidence for a stage-of-cycle effect in the predicted direction.

It is natural to ask whether this evidence of a stage-of-cycle effect in the relationship between unemployment and the dispersion measure simply reflects some broader aspect of unemployment-rate fluctuations not captured by the model. To investigate this possibility, I reestimated the benchmark specification, adding various interaction variables involving monetary disturbances. These results appear in the second four rows of Table 4. For two of the specifications, the parameter estimates did not converge, suggesting that variations in the unemployment rate associated with monetary disturbances are highly coincident with the dating of recessions and expansions. For the other two specifications, the coefficient estimates on the monetary disturbance interaction variables have the wrong sign and are not highly significant. Thus, I find no indication in the quarterly data that the stage-of-cycle effect in the unemployment/dispersion measure relationship merely reflects some broader, unexplained characteristic of unemployment-rate fluctuations.

C. TESTS BASED ON ANNUAL DATA

With the exception of Loungani (1986), previous empirical work in the sectoral shifts literature relies exclusively on post-World War II data. In this subsection, I report findings based on annual data from 1920 to 1985. The data indicate tremendous sectoral changes in employment during the years 1920 to 1946 relative to the postwar years. Thus, if we want to investigate the aggregate effects of sectoral labor reallocation, the payoff to the inclusion of these earlier years is potentially large. Despite this promise, some new problems arise by extending the sample backward in time. The earlier data are noticeably noisier and of lower quality than the data for later years. The decreased quality of the data

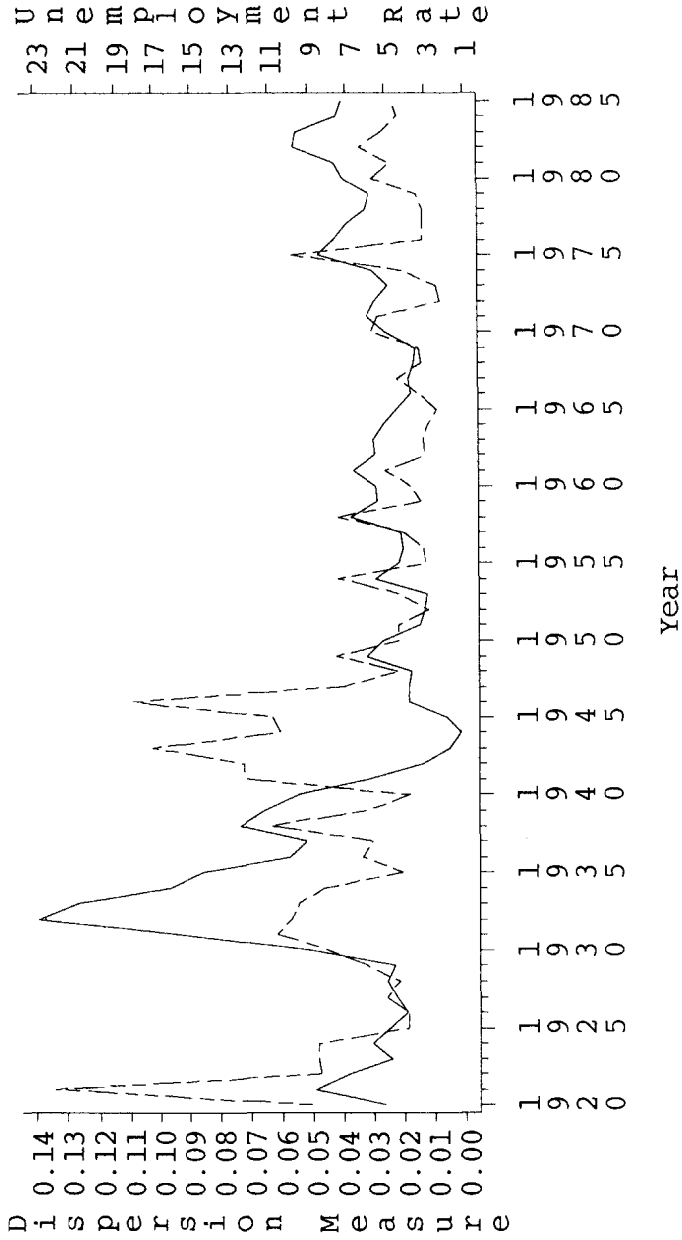
and the less-frequent sampling make it difficult to detect subtler aspects of the relationship between unemployment and other variables. Furthermore, the dramatic and violent economic fluctuations experienced by the economy during the 1930s and 1940s present new challenges, as well as new opportunities, to any empirical investigation.

Figures 6 and 7 plot the behavior of the dispersion measure σ_t , the unemployment rate, and the horizon-three covariance measure, $\sigma_{t,3}^H$.¹⁰ These plots convey some interesting points. A striking feature of both the dispersion and covariance measures is their tremendous volatility over the 1920 to 1946 years relative to the 1947 to 1985 years. This greater volatility is especially striking for the covariance measure. It seems unlikely that greater measurement error fully accounts for the greater volatility of these measures prior to 1947. Errors in the sectoral employment levels that are serially uncorrelated and uncorrelated across sectors raise the average value of the dispersion measure but not the volatility of the measure. Similarly, this pattern of measurement error lowers the average value of the covariance measure but does not cause increased volatility. Thus, Figures 6 and 7 bear out the earlier claim of tremendous changes in the sectoral distribution of employment over the 1920 to 1946 years.

The dispersion-measure plot indicates very large structural shifts in the sectoral distribution of employment during the world War II years. This period of rapid labor reallocation is highly unusual in that it coincided with very rapid expansion of output and very low unemployment rates. Aside from the World War II years, two of the largest peaks in the pace of labor reallocation, as proxied by the dispersion measure, occurred during the depths of the Great Depression in 1932 and following World War II in 1946. But the plot of the covariance measure in figure 7 illustrates a crucial difference between these two years of rapid labor reallocation: in 1946 the sectoral changes in employment reversed the changes during the preceding war years, while in 1932 the sectoral changes in employment

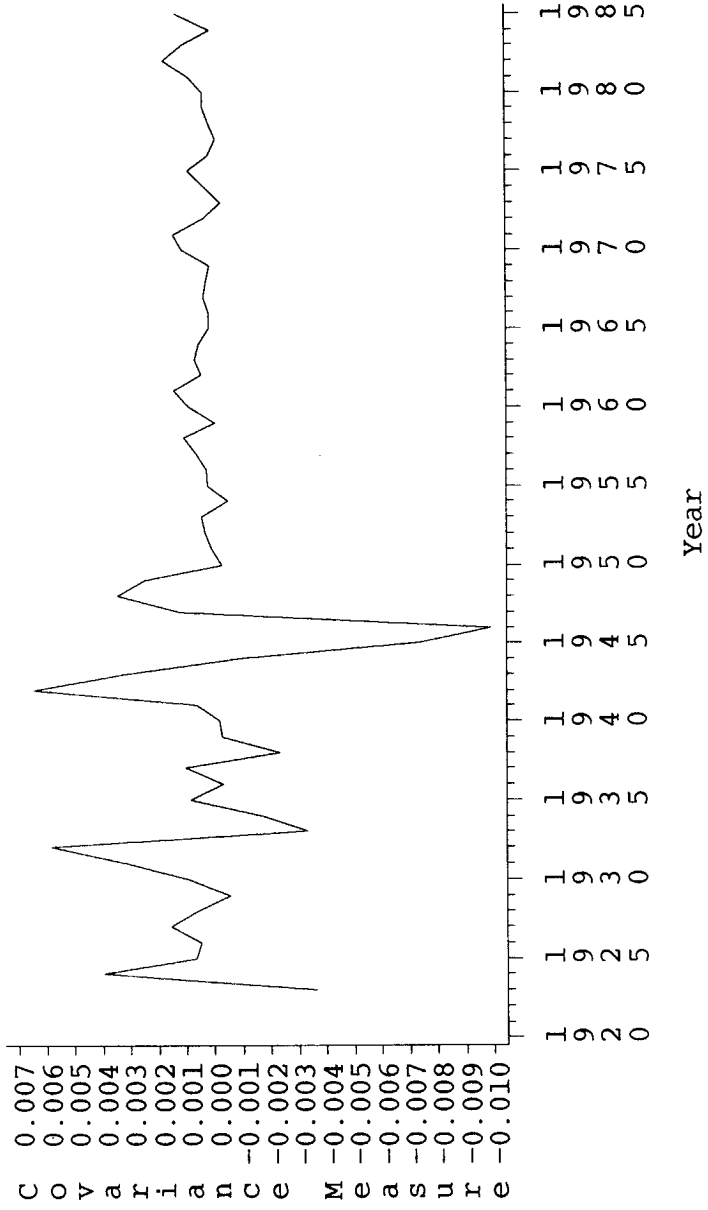
¹⁰In the annual data, the dispersion and covariance measures are based on the following eight-industry decomposition of nonagricultural employment: government, wholesale and retail trade, the FIRE group, manufacturing, transportation and public utilities contract construction, mining, and services. Note that these measures are constructed from changes in annual sectoral employment levels. The time plots for $\sigma_{t,2}^H$ and $\sigma_{t,4}^H$ are very similar to the plot for $\sigma_{t,3}^H$ in Figure 7. The plot for $\sigma_{t,1}^H$ also exhibits much more volatility prior to 1947, but its pattern of peaks and troughs differs from the pattern displayed in Figure 7.

Figure 6
Unemployment Rate and Cross-Sectoral Dispersion Measure, 1920 to 1985



The solid line represents the civilian unemployment rate.
 The dashed line represents the cross-sectoral dispersion measure.

Figure 7
Weighted Cross-Sectoral Covariance at Horizon Three, 1923 to 1985



reinforced the pattern of change during preceding years. Figure 7 also indicates that the estimated effects of the covariance measure on the unemployment rate will be heavily influenced by a few extreme values assumed by this measure during the first half of the sample.

In my formal econometric work with the annual data, I draw heavily on a recent study by Rush (1986). Rush investigates the effects of monetary disturbances and other factors on the unemployment rate over the 1920 to 1983 period. A notable feature of Rush's framework is his focus on the effects that anticipated and unanticipated movements in the monetary base have on the unemployment rate. I do not repeat Rush's specification and stability tests, nor do I investigate whether only unanticipated monetary base movements affect unemployment. I simply adopt those specifications that emerge most favorably from his study as a useful starting point for my investigation.

A typical unemployment equation estimated with the annual data is given by

$$UN_t = \theta_0 + \theta_1 TIME_t + \theta_2 DBU_t + \theta_3 DBU_{t-1} + \theta_4 DBU_{t-2} + \theta_5 DMULT_t + \theta_6 DMULT_{t-1} + \theta_7 \sigma_t + \theta_8 \sigma_{t-1} + \theta_9 \sigma_{t-2} + \theta_{10} \sigma_{t,1}^H + \theta_{11} \sigma_{t,4}^H + \theta_{12} GPUR_t + u_{1t}, \quad (5)$$

where DBU is the unanticipated component of the monetary base growth rate (computed as the change in the log), DMULT is the growth rate of the M2 money-supply multiplier, GPUR is the log of real government purchases of goods and services, u_1 is an error term, and the other variables have been previously defined. The money-multiplier variable is included in some estimated specifications as a device to control for the effects of the collapse of financial intermediation during the 1930s. See Bernanke (1983).

The forecasting equation for the base growth equation is

$$DB_t = \alpha_0 + \alpha_1 DB_{t-1} + \alpha_2 UN_{t-1} + \alpha_3 FEDVP_t + u_{2t} \quad (6)$$

where $FEDVP_t$ is the (predicted) difference between the actual growth rate of federal government expenditures and the "normal" growth rate of federal government expenditure. The inclusion of FEDVP is designed to capture the revenue motive for money creation associated with temporary changes in federal government spending (when it is costly to vary the amount of revenue raised from other sources). $FEDVP_t$ is based entirely on information

available at time t-1 and is calculated according to

$$FEDVP_t = .2(\hat{DG}_t + \sum_{i=1}^{15} .8^i DG_{t-i}),$$

where DG_t is the growth rate in federal government expenditure at t, and \hat{DG}_t is the predicted growth rate, based on period t-1 information. \hat{DG}_t is the predicted value from the following equation:

$$DG_t = \lambda_0 + \lambda_1 DG_{t-1} + \lambda_2 UN_{t-1} + \lambda_3 WAR_t. \quad (6)$$

WAR_t equals zero, except in a year that immediately follows the end of a war. In these years, WAR_t equals the annual average of the number who served in the military during the preceding war.

Aside from the addition of the dispersion and covariance measures in the unemployment equation, my estimated equations differ somewhat from Rush's specifications. First, most of the results I report are based on a 1924 to 1985 sample period. In my data set, 1924 is the first year for which $\sigma_{t,4}^H$ can be calculated. Rush reports results for the 1920 to 1983 and 1918 to 1983 sample periods. Second, Rush includes a measure of the proportion of total banks that suspended operations during the year in the base growth forecasting equation. This variable never showed a significant effect in my work, so I dropped it from the reported estimates. Third, to maintain comparability with my results based on quarterly data, I report results based on an untransformed unemployment-rate variables. (Transforming the unemployment-rate variable as in Rush does not alter my findings). Fourth, my data sources differ from Rush's sources for several variables, most notably the unemployment rate. See the appendix for details. Finally, my results are based on joint estimation of the unemployment equation and two forecasting equations.

Table 5 reports the results of estimating different variants of the model (4)-(6) over various sample periods. In discussing Table 5, I confine myself primarily to the estimated effects of the dispersion and covariance measures on the unemployment rate. Other aspects of Table 5 are largely consistent with results in Rush (1986).

Column (1) of Table 5 reports results of estimating the three-equation system over the 1924 to 1985 sample. The unemployment equation specification in column (1) contains the dispersion measure and its first two lags, covariance measures at horizon one and horizon four, and the

variables included in Rush's study. The dispersion-measure coefficient estimate has a marginal significance level of less than .02 and implies quantitatively large effects on the unemployment rate. The point estimate of 3.54, and the sample range of movements in the dispersion measure, imply a range of variation in the unemployment rate of 3.60 percentage points. The point estimates on the lagged values of the dispersion measure add another 2.29 percentage points, but these point estimates do not differ from zero at conventional significance levels. Most of the other estimated specifications, reported in columns (2)-(7), imply even larger effects of the dispersion measure on the unemployment rate. Thus, the estimated effects of the dispersion measure on the unemployment rate in Table 5 echo the results in Table 1 based on postwar quarterly data. Table 5, however, provides weaker evidence that lagged values of the dispersion measure affect unemployment.

The most interesting finding in Table 5 is the strong positive effect of long-horizon covariance measures on the unemployment rate. In column (1) the coefficient estimate on $\sigma_{t,4}^H$ is large, positive, and has a marginal significance level of .001. The point estimate, and the sample range of variation in $\sigma_{t,4}^H$, imply a range of variation in the unemployment rate of 6.00 percentage points. Results based on $\sigma_{t,3}^H$ are similar (see column (3)). Recalling Figure 7, these results indicate that the long-horizon covariance measures account for substantial unemployment-rate movements over the pre-1947 period but much milder movements over the 1947-1985 period.

The finding that long-horizon covariance measures account for substantial movements in the unemployment rate carries important implications. Most directly, this finding supports the view that past patterns of sectoral labor reallocation have important effects on current unemployment behavior. If past patterns of labor reallocation have important effects on current unemployment, then variations in the current pace of labor reallocation will have important effects on unemployment. Hence, indirectly, this finding lends credence to the view that the positive correlation between cross-sectoral employment-dispersion measures and the unemployment rate reflects the role of fluctuations in the pace of labor reallocation. This finding also provides a partial explanation for an observation that has long puzzled economists: unemployment-rate fluctuations exhibit great persistence. If the patterns of labor reallocation over several preceding years, as well as past magnitudes of labor reallocation, affect current unemployment behavior, then it is not

Table 5

Joint Nonlinear Least Squares Estimation of the Unemployment and Forecasting Equations with Annual Data

Specification	(1) 1924-1985	(2) 1924-1985	(3) 1924-1985	(4) 1924-1985	(5) 1953-1985	(6) 1924-1985	(7) 1925-1985
<u>Unemployment Equation</u>							
C	-9.73 (.000)	-10.82 (.011)	-11.21 (.013)	-13.92 (.003)	-5.31 (.097)	-11.36 (.000)	-5.02 (.202)
DBU	-29.7 (.001)	-25.8 (.003)	-26.5 (.005)	-43.8 (.002)	-18.5 (.002)	-29.4 (.001)	-20.2 (.015)
DBU ₋₁	-30.0 (.000)	-34.3 (.000)	-34.3 (.000)	-28.1 (.025)	-15.2 (.008)	-31.4 (.000)	-19.9 (.015)
DBU ₋₂	-29.7 (.007)	-23.3 (.003)	-22.3 (.009)	-32.8 (.011)	-12.7 (.020)	-17.9 (.011)	-12.6 (.062)
TIME	.283 (.000)	.303 (.000)	.319 (.000)	.334 (.000)	.145 (.028)	.304 (.000)	.186 (.005)
GPUR	-4.99 (.000)	-5.47 (.002)	-5.79 (.002)	-6.79 (.000)	-.65 (.699)	-5.34 (.000)	-3.39 (.002)
DMULT	-33.1 (.000)	-29.1 (.000)	-31.6 (.000)			-29.9 (.000)	-25.0 (.000)
DMULT ₋₁	-22.1 (.001)	-23.6 (.001)	-25.7 (.001)			-20.0 (.004)	-20.4 (.002)
σ_0	3.54 (.017)	4.24 (.016)	3.57 (.005)	6.06 (.031)	5.28 (.001)	3.82 (.012)	3.22 (.028)
σ_{-1}	1.41 (.294)	1.20 (.435)	.804 (.629)	1.48 (.564)	2.14 (.066)	2.22 (.107)	1.99 (.132)
σ_{-2}	.85 (.485)	-.087 (.951)	-.754 (.613)	2.31 (.322)	-.17 (.844)	1.06 (.396)	1.58 (.196)
σ_1^H	-7.9 (.724)	7.96 (.738)	6.13 (8.11)	-24.67 (.566)	-8.6 (.905)	-13.8 (.597)	2.15 (.917)
H			44.18 (.005)				
σ_5^H							
σ_4^H	28.9 (.001)	42.15 (.000)		37.04 (.016)	24.74 (.239)	32.09 (.001)	19.51 (.016)
Y41		.957 (.650)	1.31 (.566)				
Y42		-1.55 (.628)	-1.24 (.725)				
Y43		-2.15 (.574)	.349 (.929)				
Y44		1.55 (.660)	3.06 (.415)				
Y45		4.99 (.110)	5.56 (.098)				
AR(1) parameter							.749 (.000)
Pre-1952 weighting?	no	no	no	no	no	yes	no
R ²	.906	.921	.908	.624	.921	.876	.914
S.E.E.	1.48	1.44	1.54	2.89	0.60	1.16	1.44
DW	1.26	1.39	1.29	0.82	1.14	1.18	1.67

Table 5 continued

	Base Growth Forecasting Equation						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
C	-.019 (.008)	-.020 (.007)	-.018 (.016)	-.026 (.001)	-.026 (.080)	-.021 (.003)	-.007 (.427)
UN ₋₁	.0044 (.000)	.0047 (.000)	.0046 (.000)	.0054 (.000)	.0080 (.001)	.0047 (.000)	.0030 (.003)
FEDV P	.378 (.000)	.356 (.000)	.343 (.000)	.356 (.000)	.143 (.392)	.366 (.000)	.310 (.000)
DB ₋₁	.441 (.000)	.453 (.000)	.442 (.000)	.474 (.000)	.444 (.001)	.474 (.000)	.454 (.000)
Pre-1952 weighting?	no	no	no	no	no	yes	no
R ²	.646	.644	.653	.621	.441	.636	.685
S.E.E.	.0324	.0325	.0321	.0335	.0270	.0238	.0308
DW	1.65	1.65	1.67	1.58	1.79	1.79	1.84
Federal Expenditure Growth Forecasting Equation							
C	.029 (.407)	.028 (.417)	.028 (.416)	.027 (.435)	.116 (.002)	.021 (.486)	.031 (.383)
WAR	-.142 (.000)	-.141 (.000)	-.141 (.000)	-.137 (.000)	-.158 (.002)	-.140 (.000)	-.137 (.000)
UN ₋₁	.0043 (.324)	.0043 (.328)	.0043 (.326)	.0045 (.303)	-.0110 (.062)	.0038 (.360)	.0043 (.323)
DG ₋₁	.396 (.000)	.408 (.000)	.406 (.000)	.393 (.000)	.069 (.526)	.410 (.000)	.385 (.000)
Pre-1952 weighting?	no	no	no	no	no	yes	no
R ²	.578	.578	.578	.579	.243	.571	.581
S.E.E.	.140	.140	.140	.140	.058	.091	.140
DW	1.76	1.77	1.77	1.74	1.95	1.81	1.74

Notes: i. Table entries provide coefficient estimates for the indicated variables. The approximate marginal significance level on a two-tailed t-test against the null hypothesis of a zero coefficient appears in parentheses.

ii. The marginal significance level is based on a degrees of freedom correction that attributes 1/NUM degrees of freedom to each equation for the shared parameters in the model, where NUM equals the number of equations that share the parameter.²

iii. S.E.E. is the root mean square error of the estimated equation, DW is the Durbin-Watson statistic. R² is computed as one minus the quantity, error sum of squares divided by the dependent variable corrected mean sum of squares.

iv. The three equations are estimated jointly by nonlinear least squares, imposing the cross-equation restrictions implied by the forecasting mechanisms. The variance of the disturbance term is allowed to vary across equations. Except for column (6), the disturbance term is assumed to be homoscedastic within each equation. In column (6), pre-1952 observations are weighted by .4.

v. The dispersion and covariance measures are scaled up by a factor of 10 prior to estimation.

vi. Y41, Y42, etc. are year dummies. These variables equal one in the indicated year and zero otherwise.

ii. The specification in column (7) is estimated with a first-order autoregressive error correction in the unemployment equation.

surprising that unemployment-rate fluctuations exhibit great persistence. Of course, it remains to explain exactly why past patterns and magnitudes of labor reallocation over long horizons affect current unemployment behavior. I have suggested several reasons for these effects, but a fuller exploration of this question lies beyond the scope of this paper.

In contrast to the large effects of the long-horizon covariance measures, Table 5 provides no evidence that the short-horizon covariance measure, $\sigma_{t,1}^H$, accounts for unemployment movements. This result suggests that short-horizon covariance measures do not capture the effects of important changes in the sectoral allocation of labor resources. A pattern of labor reallocation that persists over one year causes less significant changes in the sectoral allocation of labor resources than a pattern of labor reallocation that persists over three or four years.

Columns (2)-(7) in Table 5 indicate that effects of the dispersion and covariance measures on the unemployment rate are insensitive to a number of specification changes. In column (2) separate dummies for each year of World War II are added to the unemployment equation. The coefficient estimates on the year dummies are not highly significant, nor are the other estimates greatly affected by inclusion of the year dummies. According to the estimated model, sharp increases in government purchases more than offset the unemployment effects of unusually rapid labor reallocation during the war years. Column (4) in Table 5 indicates that the point estimate on $\sigma_{t,4}^H$ from a 1953 to 1985 sample is approximately the same size as the point estimate from the 1924 to 1985 sample. The estimated standard error rises sharply, however, so that the marginal significance level in the postwar sample is .25. Column (5) indicates that the estimated effects of the dispersion and covariance measures are not greatly affected by the exclusion of the money-supply multiplier variables. Inspection of residual plots suggests a larger error-term variance in all three equations over the 1924-1951 subsample than over the 1952-1985 subsample. To correct for this heteroscedasticity, column (6) reports results based on a scheme that weights the pre-1952 observations 40% as heavily as the remaining observations. As compared to column (1), these results indicate stronger evidence that lagged values of the dispersion measure affect unemployment. The Durbin-Watson statistic and the residual plots also indicate positive serial correlation in the unemployment equation. Column (7) reports results based on a first-order autoregressive error correction in the unemployment equation. These results, too, provide stronger evidence that lagged values of the dispersion measure affect

unemployment. Finally, I have estimated models with an additional lag of DB in the base-growth equation and an additional lag of DG in the expenditure-growth equation. These specification changes do not affect the results.

Turning to the stage-of-business-cycle effects, tests based on annual data (not shown in Table 5) yield no evidence that the unemployment/dispersion measure relationship varies systematically over the cycle. Point estimates on the recession/dispersion measure interaction variables are positive, as predicted, but the estimated standard errors are so large that no inference can be drawn. In some cases, I did not obtain convergent parameter estimates.

IV. THE TIMING OF LABOR REALLOCATION

According to the reallocation-timing hypothesis, labor mobility and turnover are highly substitutable over time. Coupled with the observation that labor mobility and turnover involve unemployment and other forms of foregone production, this statement carries a fundamental implication: movements in the value of foregone production induce negatively correlated movements in the pace of labor reallocation and unemployment. At the aggregate level, this implication translates into the testable prediction that the cross-sectoral average value of foregone production is negatively correlated with the amount of unemployment due to labor reallocation. I perform simple tests of this prediction about reallocation timing.

To test this prediction, one needs a suitable proxy for the cross-sectoral average value of foregone production. I use two real wage measures, two inventory measures, and an inventory/sales ratio to proxy for the value of foregone production. The motivation for the use of real wage measures is clear, but some points require elaboration. First, the average hourly wage inclusive of overtime is the appropriate measure. Employers' willingness to pay overtime premiums indicates they attach a high value to foregone production. Second, a broad-based wage measure is desirable, since the measure must proxy for the average value of foregone production across sectors. I use an index of hourly compensation for wage and salary workers in the nonfarm business sector. This index is based on estimated hours worked. Third, the timing of labor reallocation reflects the outcome of a joint maximization problem confronting workers and firms. Whether the labor-demand price (wage deflated by output price) or labor-supply price

(wage deflated by cost of living) or an average of the two best proxies for the value of foregone production is unclear. Hence, I deflate by the producer price index in one wage measure and by the consumer price index in a second wage measure.

As proxies for fluctuations in the value of foregone production, these wage measures suffer several oft-noted deficiencies. First, under long-term contractual arrangements between workers and firms, the real wage need not mimic short-run movements in the value of labor's marginal product. Second, (effective) hours worked--hence, hourly wages--are notoriously difficult to estimate for many workers. Wage-measurement error that fluctuates systematically over the business cycle can bias the tests below. Third, selectivity phenomena also impart a cyclically varying bias to observed aggregate wages. Bils (1985), Heckman and Sedlacek (1985), and Moffitt, Runkle, and Kean (1986) discuss the aggregation bias that arises from selectivity phenomena and investigate the magnitude of this bias in cyclic real-wage movements.

These problems with the wage measures argue for the use of other types of proxies for the average value of foregone production. The two beginning-of-period inventory measures I use are the book value of manufacturers' finished-goods inventory, deflated by the producer price index, and a constant dollar manufacturers' finished-goods inventory series based on separate, industry-specific price deflators and adjustments for LIFO and non-LIFO inventory valuation methods. The constant dollar inventory series is the more appropriate measure, but the book value series is available over a longer time period.

Finished-goods inventories serve as a buffer stock that enables firms to economize on costs of adjusting production and employment in the face of fluctuating output demand.¹¹ An unanticipated decrease in demand for the firm's output--i.e., a fall in the value of its product--leads to inventory accumulation. If the demand decrease persists, inventory carrying costs grow as inventories continue to accumulate, prompting the firm to scale back production and decrease employment. In this scenario, unusually high finished-goods inventory levels signal a lower value of foregone production and trigger layoffs. Similarly, unusually low inventory levels signal a higher value of foregone production. Thus, according to this scenario, the

¹¹ I adopt the usual assumption that inventory fluctuations are demand-driven.

reallocation-timing hypothesis predicts a positive correlation between detrended finished goods inventories and proxies for the quantity of labor reallocation.

Inventory levels also fluctuate in response to anticipated future demand disturbances. For example, firms can accumulate inventories in anticipation of higher future demand. Thus, high (low) inventory levels need not always signal a low (high) value of foregone production. The use of seasonally adjusted data removes the most important component of predictable demand fluctuations, but the inventory measures are still poor proxies for the value of production to the extent that the timing and magnitude of nonseasonal demand fluctuations are predictable. To address this problem, I also use an inventory/sales ratio--the constant dollar finished-goods inventory series divided by a constant dollar series on manufacturers' sales. Unanticipated demand disturbances cause inventories and sales to move in opposite directions, whereas anticipated demand disturbances cause inventories and sales to move in the same direction. (Of course, the timing of movements in inventories and sales differs in response to anticipated disturbances.)

As proxies for the pace of labor reallocation, I use the simulated unemployment-rate series plotted in figure 5 and the employment growth dispersion measure described by equation (1). For purposes of comparison, I also include a variable that measures the actual unemployment rate, net of the simulated unemployment-rate series. The real wage and inventory series were first logged, then linearly detrended. Plots of these variables revealed sharp breaks in the two real wage series and the deflated book value inventory series around 1972/1983. The constant dollar inventory plot suggested a slight break in trend behavior around this date. Hence, I detrended each of these four variables separately over the pre-1973 and post-1972 sample periods.

Table 6 reports the contemporaneous correlations between the labor reallocation proxies and the five proxies for the average value of foregone production. Separate results for the full sample as well as the pre-1973 and post-1972 subsamples appear in Table 6. These results bear out the prediction of the reallocation-timing hypothesis. For the full samples, the correlations between the two labor reallocation proxies and the five average value proxies have the predicted sign in every case--in every case but one, the correlation differs from zero at the five-percent significance level. Furthermore, UN - UNSIM, the unemployment rate net of the simulated unemployment rate, shows evidence of a nonzero correlation only with the

Table 6

Contemporaneous Correlations Between Labor Reallocation
Measures and Average Value Proxies

	Detrended		Detrended	
	$\log\left(\frac{\text{Compensation Index}}{\text{Producer Price Index}}\right)$		$\log\left(\frac{\text{Compensation Index}}{\text{Consumer Price Index}}\right)$	
Sample	1953:2 - 1986:2		1953:2 - 1986:2	
UNSIM	-.227	(.009)	-.306	(.000)
σ	-.217	(.012)	-.250	(.004)
UN - UNSIM	-.085	(.328)	.013	(.881)
Sample	1953:2 - 1972:4		1953:2 - 1972:4	
UNSIM	-.085	(.454)	-.376	(.001)
σ	-.127	(.264)	-.351	(.002)
UN - UNSIM	-.129	(.257)	.129	(.257)
Sample	1973:1 - 1986:2		1973:1 - 1986:2	
UNSIM	-.454	(.001)	-.234	(.089)
σ	-.257	(.061)	-.180	(.193)
UN - UNSIM	-.149	(.282)	-.113	(.417)

Table 6 continued

Contemporaneous Correlations Between Labor Reallocation Measures
and Average Value Proxies

Detrended		Detrended		
Mfg's Fin Goods $\log\left(\frac{\text{Inv'y, Book Value}}{\text{Producer Price Index}}\right)$		Mfg's Fin Goods $\log\left(\frac{\text{Inv'y, Constant \$}}{\text{Inv'y, Constant \$}}\right)$		$\log\left(\frac{\text{Constant \$ Inv'y}}{\text{Mfgs' Sales}}\right)$
1953:2 - 1986:2		1959:2 - 1986:1		1961:1 - 1985:4
.091 (.298)		.212 (.028)		.710 (.000)
.185 (.033)		.219 (.023)		.359 (.000)
-.056 (.523)		-.010 (.923)		.561 (.000)
1953:2 - 1972:4		1959:2 - 1972:4		1961:1 - 1972:4
.031 (.789)		.098 (.477)		.764 (.000)
.244 (.030)		.230 (.091)		.527 (.000)
-.180 (.113)		-.216 (.114)		.427 (.003)
1973:1 - 1986:2		1973:1 - 1986:2		1973:1 - 1985:4
.210 (.128)		.395 (.004)		.833 (.000)
.145 (.294)		.235 (.090)		.308 (.026)
.033 (.816)		.121 (.389)		.822 (.000)

Notes:

- i. Table entries are Pearson sample correlation coefficients between the row and column variables. The number in parentheses is the marginal significance level for a test against the null hypothesis that the population correlation equals zero.
- ii. UNSIM is the simulated unemployment rate series plotted in Figure 5. It is based on parameter estimates in Table 1.
- iii. Each detrended variable is the residual from a regression on a constant and linear time trend. Separate regressions were run over the pre-1973 and post-1972 sample periods for each detrended variable.

inventory/sales ratio. The same pattern of results holds in the two subsamples, except that in some cases the correlations between the reallocation proxies and average value proxies are insignificantly different from zero.

Figure 8 plots the inventory/sales ratio for the 1961:1 to 1985:4 sample period, overlaid by a plot of the simulated unemployment-rate series. Through 1977 the two plots display similar amplitudes and a remarkably similar timing. After 1977 the simulated unemployment-rate series lags noticeably behind the inventory/sales ratio.

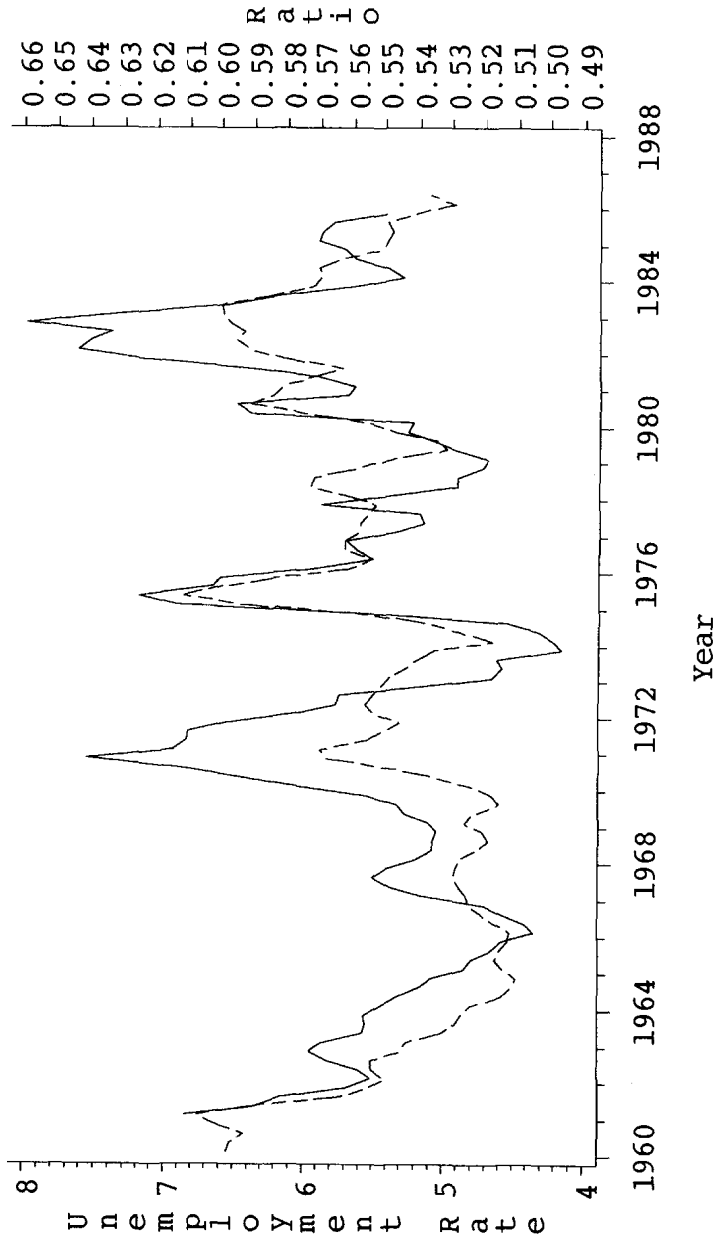
While these results confirm a fundamental prediction of the reallocation-timing hypothesis, it is possible to reconcile them with the two competing hypotheses. A wide range of cyclic real wage behavior is consistent with the normal business-cycle hypothesis--see Leiderman (1983) for a summary discussion and references to the relevant literature. Furthermore, the same intertemporal substitution effects that drive the timing of permanent separations also drive the timing of temporary separations, so that the normal business-cycle hypothesis can account for the inventory and inventory/sales ratio correlations.

The sectoral shifts hypothesis can explain the real wage correlations along the following lines (see Black (1982)). A close match between the desired and actual labor-force allocation implies that market goods will be valued highly, people will spend much time working, and the marginal value of leisure and the real wage will be high. A poor match implies that market goods will not be highly valued, so people will work less, driving down the marginal value of leisure into equality with the lower real wage.¹² The inventory and inventory/sales ratio correlations can be reconciled with the sectoral shifts hypothesis by noting that the inventory measures cover only the manufacturing sector and most large movements in the dispersion measure, over the sample, reflect declines in durable goods manufacturing employment.

Summarizing, the tests in this section confirm the basic prediction of the reallocation-timing hypothesis that the pace of labor reallocation moves inversely to the opportunity cost of labor mobility. It is possible, however, to reconcile the test results with the sectoral shifts and normal

¹²Working against this real-wage effect is the following; when the match is a bad one, the returns to labor reallocation are high, encouraging people to spend more time at switching sectors, which drives up the marginal value of leisure and the real wage.

Figure 8
Inventory/Sales Ratio and Simulated Unemployment Rate, 1961:1 to 1985:4



Solid Line: ratio, constant dollar inventories to sales, manufacturing.
 Dashed Line: simulated unemployment rate from Table 1.

business-cycle hypotheses.

V. AN ASSESSMENT OF THE THREE COMPETING HYPOTHESES AND CONCLUDING REMARKS

With respect to the competing hypotheses about unemployment-rate fluctuations, several messages emerge from Sections II-IV:

(1) The strong form of the normal business-cycle hypothesis is decisively rejected. Put more positively, fluctuations in the pace of labor reallocation are a key aspect of short-run unemployment-rate fluctuations. The cyclic behavior of unemployment inflows and outflows, coupled with the relative magnitudes of permanent and temporary separations, provides direct evidence that fluctuations in the pace of labor reallocation contribute greatly to unemployment fluctuations. Stated differently, a large fraction of unemployment-rate fluctuations reflects fluctuations in frictional and structural unemployment.

(2) Long-horizon cross-sectoral covariance measures and long lags of cross-sectoral dispersion measures greatly affect unemployment. These findings indicate that past patterns and magnitudes of labor reallocation strongly influence current unemployment behavior, confirming key predictions of the sectoral shifts hypothesis and providing further support for the view that fluctuations in the pace of labor reallocation constitute a key feature of unemployment-rate fluctuations.

(3) Past labor reallocation can affect current unemployment behavior in two distinct ways. First, by interrupting long-term firm-worker matches, past allocative disturbances induce workers to repeat the matching and turnover process that is typical of new labor-force entrants. Thus, allocative disturbances increase the unemployment propensities of affected workers for several years. Second, past allocative disturbances determine the current degree of skill, locational, and informational mismatch between firms and workers. Thus, the contemporaneous response to an allocative disturbance depends on its relationship to past patterns of labor reallocation. These observations and the corroborating empirical findings provide a partial explanation for the great persistence of unemployment-rate fluctuations.

(4) The prediction of the sectoral shifts and reallocation-timing hypotheses that a rapid rate of attachments to new jobs accompanies cyclically high unemployment rates is fully consistent with the observed

behavior of vacancy stock measures. The cyclic behavior of unemployment inflows and outflows indicates that periods of high unemployment rates and low vacancy stocks are also periods of high flow rates of new vacancies.

(5) The basic prediction of the reallocation-timing hypothesis that the pace of labor reallocation moves inversely to the opportunity cost of labor mobility finds support using several different proxies for the cross-sectoral average value of foregone production.

Are these messages consistent with other evidence in the sectoral shifts literature?

My earlier paper, Davis (1986b), documents the temporally asymmetric character of unemployment-rate fluctuations associated with movements in the cross-sectoral dispersion measure. The unemployment-rate fluctuations are temporally asymmetric in the sense that, on average, increases in unemployment are sharper and of shorter duration than decreases. For example, using the methods described in my earlier paper, the simulated unemployment-rate series in Figure 5 implies an estimated unconditional duration of increases that is only 74% as long as the estimated unconditional duration of decreases.¹³ This temporal asymmetry finding confirms a prediction of the sectoral-shifts hypothesis. The reallocation-timing hypothesis also suggests an explanation for this finding. A transitory decline in the average value of production triggers an increase in the job-separation rate and the pace of labor reallocation, leading to an increase in the unemployment rate. Because the worker-employer match formation process is more time-consuming than the match breakup process, unemployment declines will be "stretched out" over time relative to unemployment increases. (See Pissarides (1985)).

Other evidence supports the sectoral shifts hypothesis and is difficult to reconcile with either of the other two hypotheses about unemployment-rate fluctuations. The sectoral shifts hypothesis carries the distinctive implication that the quantity of labor reallocation depends on

¹³This 74% figure is similar to results obtained under a variety of different specifications in my earlier work. By way of contrast, I find no evidence in my earlier work that monetary disturbances cause temporally asymmetric unemployment-rate fluctuations.

the magnitude, and not the direction, of allocative disturbances.¹⁴ Under the reallocation-timing and normal business-cycle hypotheses, both the magnitude and direction of disturbances affect unemployment and, in the case of the reallocation-timing hypothesis, the quantity of labor reallocation. To test this distinctive implication of the sectoral shifts hypothesis, Davis (1986b) investigates whether the magnitude of oil price disturbances and the cross-sectoral dispersion measure exhibit a positive relationship when controlling for variables that reflect the magnitude and direction of oil price disturbances. The evidence on this question is mixed but, on balance, favorable to the sectoral shifts hypothesis. The evidence clearly indicates a strong empirical link between oil price shocks and the cross-sectoral dispersion measure. Loungani (1986) shows that when the relative price of oil is held fixed, Lilien-type dispersion measures have no remaining explanatory power for unemployment in the postwar period or in the 1900-1929 period.

Another finding that is difficult to reconcile with the reallocation-timing hypothesis is the absence of evidence that monetary disturbances influence the timing of labor reallocation. Using annual data, Lilien (1982) reports that unanticipated monetary disturbances are virtually orthogonal to the cross-sectoral dispersion measure. Using quarterly data, I find neither evidence that unanticipated monetary disturbances Granger cause the dispersion measure nor evidence that unanticipated monetary disturbances are contemporaneously correlated with the dispersion measure. I find only weak evidence that anticipated monetary disturbances are positively related to the dispersion measure.¹⁵ Furthermore, when the dispersion measure is constructed from sectoral employment regression residuals that are orthogonal to current and lagged monetary disturbances, the measure continues to show substantial explanatory power in unemployment-rate regressions--see Lilien (1983), Davis (1986b), and Loungani (1986). These findings are puzzling, because we customarily think of monetary disturbances as the most important source of aggregate-demand disturbances, and because aggregate-demand disturbances affect the average

¹⁴This statement refers to the unconditional relationship between allocative disturbances and the quantity of labor reallocation. It does not contradict earlier statements about the importance of direction in the conditional relationship between allocative disturbances and the quantity of labor reallocation.

¹⁵This evidence is reported in early versions of my (1986b) paper.

value of production. These findings suggest that, however monetary disturbances affect the aggregate economy, it is not by influencing the pace of labor reallocation.

Finally, Hamilton's (1983, 1985) striking evidence on the effects of oil price shocks on aggregate economic activity remains difficult to explain within the framework of models that abstract from specialized human and physical capital. That oil price shocks have effects in models that abstract from specialization is not at issue. The issue is how to explain the apparent magnitude of the responses to oil price shocks, particularly during the pre-OPEC period, using well-articulated models. Hamilton (1986) analyses a multisector model with specialized labor that shows how seemingly small allocative disturbances can cause large output and unemployment responses.

Taken as a whole, the evidence in this and other papers provides convincing support for the view that short-run unemployment rate fluctuations partly, perhaps largely, reflect fluctuations in the pace of labor reallocation. This empirical evidence on the role of specialized resource reallocation in unemployment fluctuations is an important result that stands as a challenge to existing theories of aggregate economic fluctuations. The emphasis on simple labor-leisure tradeoffs in these theories is much overdrawn. Building and analyzing macroeconomic models that explicitly incorporate processes by which specialized resources are reallocated over time, and which allow the pace of reallocation to vary over time, promises to significantly enhance our understanding of aggregate economic fluctuations. See Davis (1987) for a development of this argument. Fortunately, to this end we can draw on a large and growing body of theoretical and empirical work in the search, matching, and specific human capital investment literatures.

APPENDIX:
DESCRIPTION OF DATA AND SOURCES

QUARTERLY DATA

Quarterly data are seasonally adjusted unless noted otherwise. For data obtained from Citibase tapes, the Citibase tape code appears in parentheses. For data obtained from Business Conditions Digest, the Business Conditions Digest series number appears in parentheses.

UN-the civilian unemployment rate, from Citibase tapes (LHUR) and recent issues of Business conditions Digest (43), quarterly averages of monthly data.

n-the civilian labor force, sum of UN, and civilian employment from Business Conditions Digest (441).

s-number of unemployed persons; 1948-1981 from Labor force Statistics, volume II; 1982-1983 from Labor Force Statistics, volume II, supplement; 1984-1986 from recent issues of Employment and Earnings.

s⁰⁻⁴-unemployed persons by duration of unemployment, less than five weeks; same sources as for s.

ϕ and ω -unemployment inflow and outflow rates computed from n, s, and s⁰⁻⁴ as described in the text, quarterly averages of monthly flows.

Labor Force Participation Rate-the population of working age was first calculated from civilian employment and the employment-to-population-of-working-age ratio in Business Conditions Digest (90); civilian employment, unemployment, and population of working age were then used to calculate the labor force participation rate; quarterly averages of monthly data.

Normalized Help-Wanted Index-The Conference Board's Help-Wanted Index from Citibase tapes (LHEL) and recent issues of Business Conditions Digest (46), divided by civilian employment; quarterly averages of monthly data.

σ and σ_j^H -calculated from BLS establishment data on industry employment as described in the text; industry employment data from Citibase tapes (LPMI, LPCC, LPED, LPEN, LPTU, LPT, LPTR, LPFR, LPS, LPGOVF, LPGOVS) and recent issues of Employment and Earnings.

TBILL-90 day Treasury Bill rates, auction average, new issues, from citibase tapes (FYGN3) and recent issues of the Federal Reserve Bulletin, not seasonally adjusted.

RECESS-based on charts in the Handbook of Cyclical Indicators, 1984.

M1-M1 measure of the money supply; 1959-1986 from citibase tapes (FM1) and recent issues of the Federal Reserve Bulletin; earlier data from Banking and Monetary Statistics and multiplied by .992136 based on the 1959 overlap; quarterly averages of monthly data.

Producer Price Index-all commodities index, from Citibase tapes (PW) and Business Conditions Digest (330), quarterly averages of monthly data.

Consumer Price Index-all urban consumers index from the Handbook of Cyclical Indicators, 1984 and recent issues of Business Conditions Digest (320); quarterly averages of monthly data.

Index of Average Hourly Compensation, All Employees, Nonfarm Business Sector-from the 12/84 and 7/86 issues of Business Conditions Digest (345).

Manufacturers' Inventories, Finished Goods, Book Value-from 6/85 and 8/86 issues of Business Conditions Digest (65); these data are end-of-period values that were lagged one period to obtain the beginning-of-period values used in the text.

Constant Dollar Manufacturing Finished Goods Inventories-1959 to 1975 from Hinrichs and Eckman (1981), multiplied by 2.234 based on 1976 overlap with the following data; 1976 to 1985 from the Survey of Current Business, February 1986; these data are end-of-period values that were lagged one period to obtain the beginning-of-period values used in the text.

Constant Dollar Manufacturing Sales-1961-1975 from Business Statistics, p. 155, multiplied by 2.409 based on 1976 overlap with the following data; 1976-1985 from the Survey of Current Business, February 1986.

ANNUAL DATA

Number of Persons Who Report Themselves as Job Leavers-unemployed persons by reason, job leavers, from the Handbook of Labor Statistics, 1985 and recent issues of Employment and Earnings.

UN-civilian unemployment rate; 1920 to 1930 from Romer (1986), Table 9, column headed "U"; 1931-1943 from Darby (1976), Table 3, column 16; 1944-1947 from Historical Statistics, series D 85-86; 1948-1985, annual averages of quarterly data described above.

Monetary Base-high powered money, millions of dollars; to 1975 from Friedman and Schwartz (1982), Table 4.8, column 10, multiplied by .980275 based on 1976 overlap with the following data; 1976-1983 from Balke and Gordon (1986), p. 786; 1984-1985 from recent issues of the Federal Reserve Bulletin.

M2-"Old M2" definition of the money supply, spliced to "New M2" definition of the money supply; to 1975 from Friedman and Schwartz (1982), Table 4.8, column 1, multiplied by 1.51239 based on 1975 overlap with the following data; 1976-1983 from Balke and Gordon (1986), p. 786; 1984-1985 from recent issues of the Federal Reserve Bulletin.

Nominal Government Purchases of Goods and Services-1919-1941 and 1947-1981 from Balke and Gordon (1986), multiplied by .98647 based on 1982 overlap with the following data; 1982-1985 from the Survey of Current Business, March 1986, NIPA table 1.1, Line 18; 1942-1946 from The National Income and Product Accounts, 1929-1976, table 1.1, multiplied by .98647 and further multiplied by .9844 based on the 1947 overlap with Balke/Gordon data.

GNP Implicit Price Deflator-to 1983 from Balke and Gordon (1986), pp. 781-783; 1984-1985 from Economic Report of the President, Table B-3, multiplied by 2.07418 based on 1983 overlap with Balke/Gordon data.

GPUR—the log of the ratio, nominal government purchases of goods and services to the GNP implicit price deflator.

Federal Government Expenditures, millions of dollars—to 1938 from Firestone (1960), multiplied by 1.00823 based on the 1939–1940 overlap with the following data; 1939–1985 from the Economic Report of the President, 1986, table B-76.

WAR—in the year following the end of a war, the yearly average of the number who had served in the war, in millions; 1.571 in 1920, 7.3 in 1946, 1.13 in 1954, .5875 in 1973, zero in all other years.

Industry Employment Data—to 1946 from the U.S. Department of Labor's Employment and Earnings, 1909–1975, spliced to annual averages of quarterly data described above based on 1947 overlap.

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