The Cyclicality of Labor Turnover: A Joint Wealth Maximizing Hypothesis
(Revised)

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Working Paper No. 86
May 1987

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Procyclical quits and countercyclical layoffs are well-known empirical regularities. If, as under the usual interpretation, quits are voluntary separations and layoffs involuntary separations, then this evidence conflicts with neoclassical models of the business cycle. Since about three-quarters of all layoffs experience intervening spells of unemployment (Gottschalk and Maloney 1985), the voluntary-involuntary interpretation of quits and layoffs suggests involuntary unemployment--in particular, countercyclical involuntary unemployment. Since disequilibrium in the labor market implies disequilibrium in at least one other market, such a result challenges the adequacy of market clearing models of consumption and investment as well.

The voluntary-involuntary interpretation has received recent theoretical support from models with asymmetric information. Contracting models under asymmetric information generate the wage rigidity required to justify the voluntary-involuntary interpretation. Hall and Lazear (1984) develop a contracting model with bilateral asymmetric information to generate procyclical quits and countercyclical layoffs. A wage bargaining variation on the bilateral asymmetric information model produces the finding for Perry and Solon (1985). Kahn (1987) models employment behavior with optimal contracting and one-sided asymmetric information to generate procyclical quits.

Wage rigidities, informational asymmetries, and inefficient separations are properties of the current explanations of the cyclicality of labor turnover. I have a simple alternative which is consistent with both the cyclical behavior of quits and layoffs and the market clearing approach: under my alternative, layoffs are not involuntary separations.

In this paper, I present a model of turnover with wage flexibility, no unresolved informational asymmetries, and efficient separations. The foundation of the model is a simple matching framework in which employment relationships form in spot markets with flexible wages. Wage flexibility generates efficient separations so the voluntary-involuntary interpretation of the quit-layoff distinction does not apply. Rather, quits are worker-initiated and layoffs firm-initiated separations; who initiates a separation depends on the pre-separation wage. The business cycle is parameterized as symmetric movements in productivity for all firms. The result is that quits and layoffs are cyclically related to aggregate output shocks, but these turnover labels are inconsequential in terms of the allocation of economic resources.

The model is consistent with a wide variety of labor market phenomena and has several refutable implications. The implications are:

(i) Total separations are countercyclical.

(ii) Quits and layoffs--but not total separations or employment--depend on lagged cyclical fluctuations, with the sign of the lagged effect opposite to that of the contemporaneous effect. In particular, the quit rate depends on cyclical shocks--that is the change in the level of the economy.

(iii) The price component of the nominal output's cyclical fluctuation has no effect on labor turnover.

(iv) A distinction between anticipated and unanticipated cyclical shocks is irrelevant.

The matching model is developed in section I. Its cyclical properties are exhibited both graphically and analytically, and several additional implications are explained. Section II contains the empirical analysis. There I document the well-known cyclical movements of quits, layoffs, and employment, and test implications (i), (ii), and (iii) on both quarterly and annual data. Also in section II, I confront a structural implication of the model with aggregated data from the Panel Study of Income Dynamics. My conclusions are presented in the final section.

The empirical evidence indicates that my specification of the joint wealth maximizing hypothesis does reasonably well in capturing the principal features of turnover and employment cyclicality. However, tests of several refutable implications are mixed. Where the results conflict with the predictions of the model, the evidence is instructive and presents quite a challenge for future research on turnover cyclicality.

I. MODEL OF CYCLICAL TURNOVER

My analysis of the cyclicality of labor turnover features two components: a model of turnover and a parameterization of the business cycle. The approach to labor turnover that I take is based on a joint wealth maximizing model (see McLaughlin 1987b). The model has two principal elements. The first is the separation decision; the second, the attachment of a quit-layoff label to a given separation. One of the essential features of the model is its matching element. In each period t, heterogeneous, risk neutral workers and firms sort into employment relationships based on the quality or output of the match. I assume matches are made in pure spot markets. In period t-1 the representative worker is matched with a firm paying a (log) wage W_{t-1}, and this worker has an opportunity (log) wage with a second firm of R_{t-1}, and a fixed nonmarket value (in logs) R^o.[1] Between periods idiosyncratic shocks,

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[1] I limit the analysis to the case of two firms. See McLaughlin (1987b) for the generalization to n firms. I also adopt two other simplifying assumptions. First, R^o is nonrandom. Randomness in R^o adds little to the analysis. Second, I assume the worker is paid his productivity value and thus captures any rents associated with the match. In McLaughlin (1987c), I

which are common knowledge, arrive changing match values. These wage offers are drawn from a bivariate density function $g_t(W_t, R_t)$ over the rectangular support $[\underline{r}_t, \bar{r}_t]^2$.[2] If W_t and R_t are both less than R^o, the worker separates to the nonmarket sector. If the opportunity wage R_t exceeds R^o and W_t, the new value of employment with the incumbent firm, the firm and worker dissolve the employment relationship, and the worker is hired by the second firm. If $R_t < W_t$ (and $W_t > R^o$), the employment relationship continues for another period at the new wage W_t. The wage rate is flexible, consequently separations are always efficient.

One specification within the joint wealth maximizing approach applies quit-layoff labels based on the pre-separation wage W_{t-1}. For simplicity, I <u>define</u> a quit (layoff) to be a separation to employment at a real wage exceeding (falling short of) the pre-separation wage W_{t-1}.[3] Separations to higher wages are labeled quits. Those to lower wages are layoffs. In addition, if R^o exceeds both W_t and R_t, the worker exits the market sector via a layoff. ($R^o < W_{t-1}$ or the worker would not have been employed at t-1.)

The model is summarized in Figure 1. To emphasize its stochastic feature, a representative iso-probability contour is depicted. If $W_t > R^o$, all draws in the half-space below the $R_t = W_t$ ray result in continued

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analyze the effect of rent sharing on turnover in the context of the current model.

[2] In general, $g_t(W_t, R_t)$ depends on the identity of the incumbent employer. For notational ease, I suppress this element.

[3] In McLaughlin (1987b), I offer a detailed analysis of the process generating the quit-layoff labels, and link initiations of wage revisions to the pre-separation wage W_{t-1}. The definition in the text is an implication in the more detailed analysis. Furthermore, such a definition has been successfully employed in the analysis of differences in quit-layoff behavior over the life cycle, by educational attainment, and by union status (McLaughlin 1987a; 1987c).

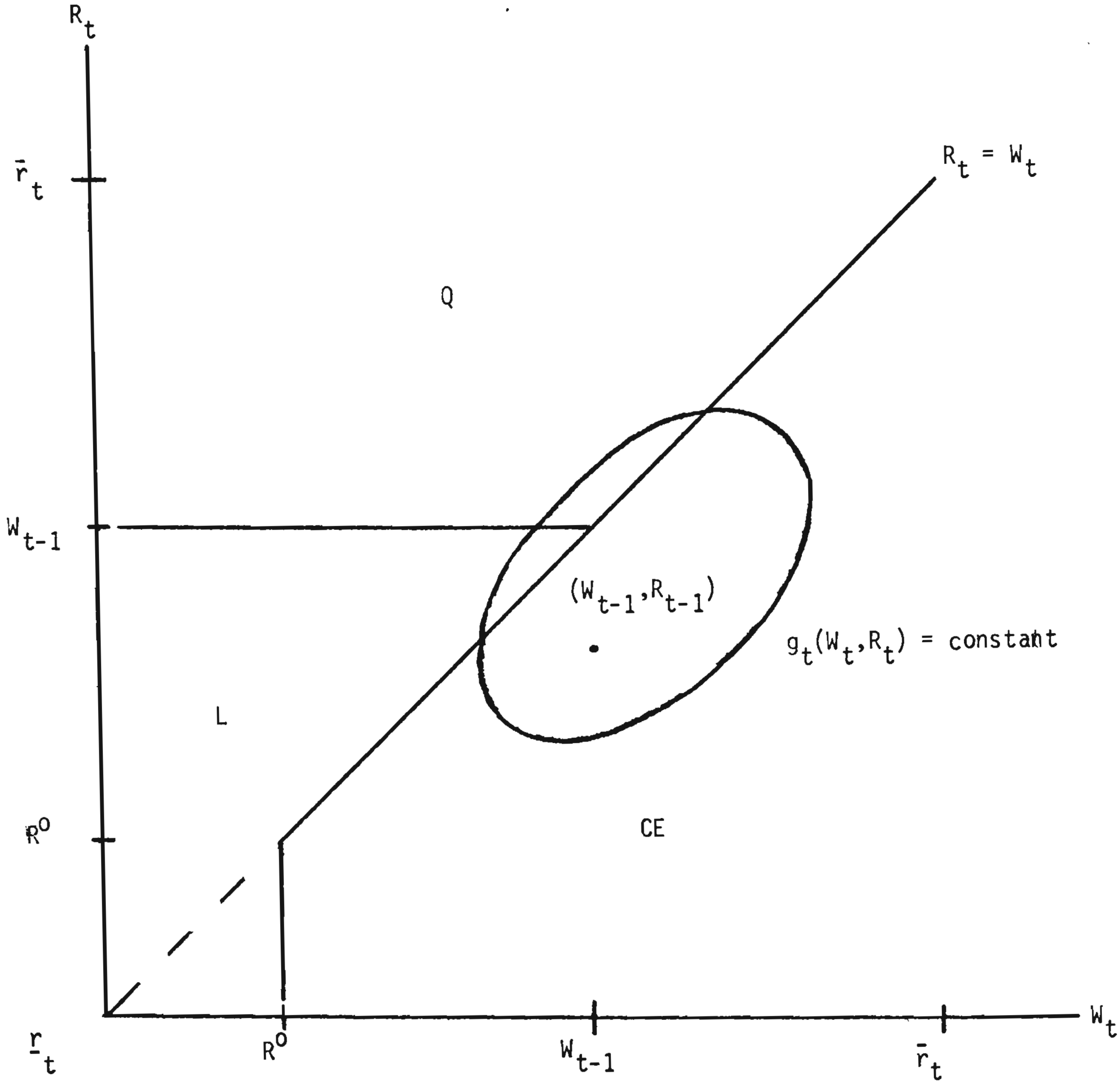

Figure 1. Turnover Regions

employment. Based on the initial draw (W_{t-1}, R_{t-1}) and the value of nonmarket production R^{o}, the quit (Q), layoff (L), and continued employment (CE) regions are defined. The turnover rates are given by the masses of the density in the Q and L regions.

The model is constructed to characterize the cyclical behavior of labor turnover. Combining the structure outlined to this point with a cyclical parameterization of the joint density which is symmetric in W_t and R_t generates the result. That is, I model cyclical growth as an equal improvement in worker productivity across the two firms. If a business cycle expansion is associated with a northeast movement of the iso-probability contour in Figure 1, then the mass rises in region Q, and falls in region L. Quits are procyclical and layoffs countercyclical.

Cyclical Structure

My analysis of turnover cyclicality requires a formal characterization of the cyclical variation in the joint density of wage offers. I assume that cyclical shocks "translate" the wage-offer distribution. Let time be indexed by τ where $\tau = 0, 1, \ldots, t, \ldots$. Consider a sequence of bivariate random variables $\{W_\tau, R_\tau\}$ which are defined over the rectangular support $[\underline{r}_\tau, \bar{r}_\tau]^2$ and correspond to the wage-offers for each τ. Denote the sequence of probability density functions associated with these random variables $\{g_\tau\}$. Business cycle shocks, which are given by a sequence $\{\xi_\tau\}$, translate the wage offer distributions in adjacent periods. That is, for a given shock ξ_τ,

$$(1.1) \qquad W_\tau = W_{\tau-1} + \xi_\tau$$

$$(1.2) \qquad R_\tau = R_{\tau-1} + \xi_\tau, \qquad \tau = 1, \ldots, t,$$

with $\xi_0 = 0$. Consequently, by repeatedly changing variables

$$(2.1)\qquad g_1(W_1, R_1) = g_0(W_1-\xi_1, R_1-\xi_1)$$

$$(2.2)\qquad g_2(W_2, R_2) = g_1(W_2-\xi_2, R_2-\xi_2) = g_0(W_2-\xi_2-\xi_1, R_2-\xi_2-\xi_1)$$

$$\vdots$$

$$(2.3)\qquad g_t(W_t, R_t) = \ldots = g_0(W_t - \sum_{\tau=1}^{t} \xi_\tau, R_t - \sum_{\tau=1}^{t} \xi_\tau) \equiv g_0(W_t-\gamma_t, R_t-\gamma_t),$$

where $\gamma_t \equiv \sum_{\tau=1}^{t} \xi_\tau$ is the deviation from the long-run level of the economy at time t. Therefore, the pair (W_t, R_t) is distributed over $[\underline{r}+\gamma_t, \bar{r}+\gamma_t]^2$ following the density function $g_0(W_t-\gamma_t, R_t-\gamma_t)$.

This representation is useful for two reasons. First, γ_t contains all the historical information required to characterize the wage-offer distribution at time t. Second, since $\gamma_t = \gamma_{t-1} + \xi_t$ and γ_{t-1} is predetermined, I can analyze the effect of ξ_t on turnover by focusing on the effect of γ_t alone.

Cyclical Implications

The cyclical implications for the turnover variables are illustrated in Figure 2 for a cyclical "boom" ($\gamma_t > 0$). The business cycle parameter translates the joint density g_0 along the ray $R_t = W_t$ by $\gamma_t' \equiv \sqrt{2}\gamma_t$ units. Outside opportunities expand by exactly the magnitude of opportunities within the incumbent firm. Since the entire density is displaced northeast by $\gamma_t' > 0$, and W_{t-1} is fixed, the mass in the quit region grows as the shaded border between L and Q moves into the Q region. Similarly, the mass in the layoff region falls as the border between Q and L moves into the Q region, and the

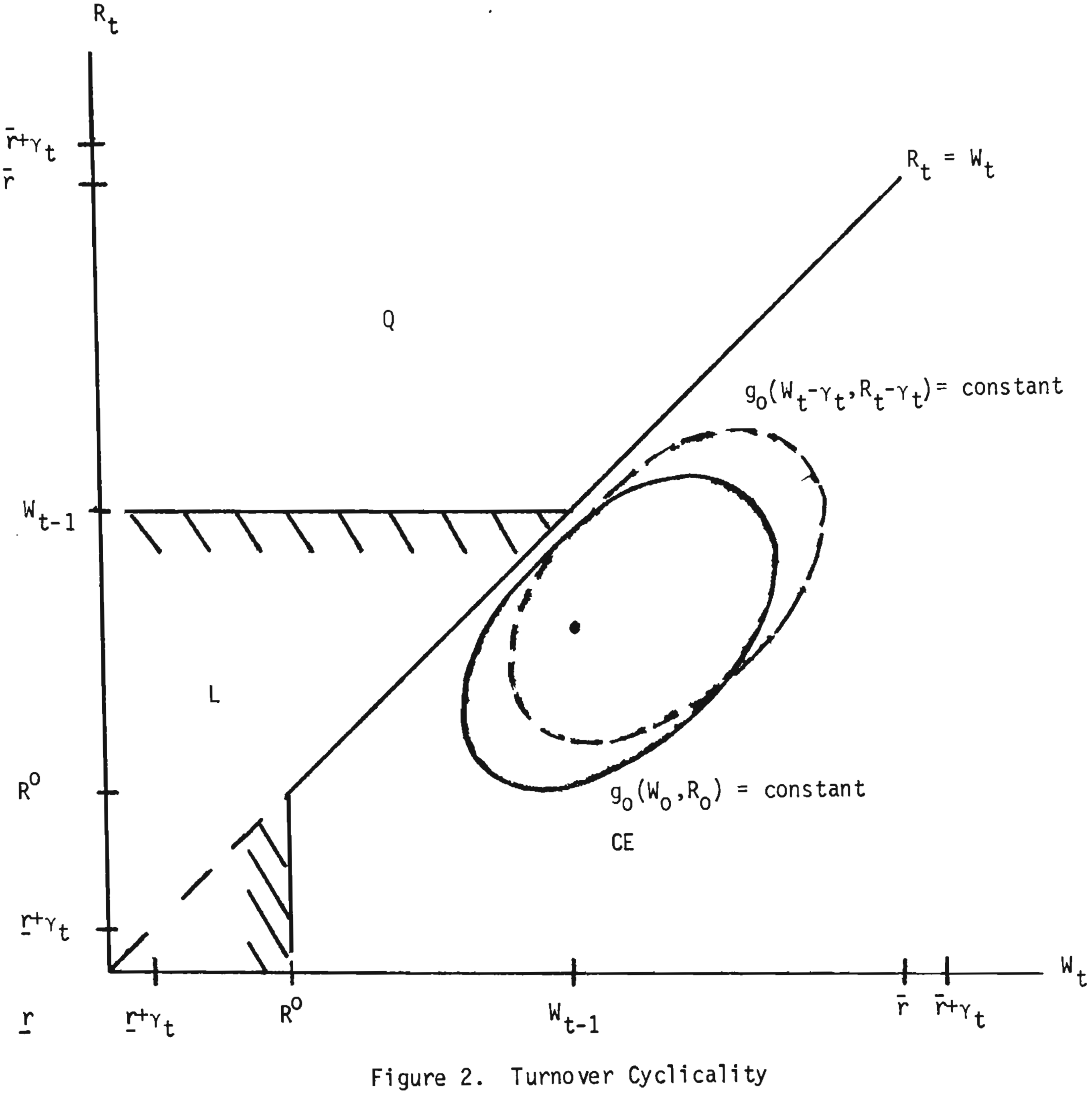

Figure 2. Turnover Cyclicality

vertical border between L and CE moves into the CE region. Quits are procyclical and layoffs countercyclical.

The turnover rates corresponding to the regional probability masses in Figure 2 have convenient analytical representations. The quit and layoff rates at time t are the probabilities associated with period t-1 to period t transitions. Conditioning on the pre-separation wage W_{t-1} and the state of the economy γ_t, the quit and layoff rates are

$$(3.1)\qquad q_t(W_{t-1}, \gamma_t) = \int_{W_{t-1}}^{\bar{r}+\gamma_t} \int_{\underline{r}+\gamma_t}^{R_t} g_o(W_t-\gamma_t, R_t-\gamma_t)\,dW_t\,dR_t$$

$$(3.2)\qquad \ell_t(W_{t-1}, \gamma_t, R^o) =$$

$$\int_{\underline{r}+\gamma_t}^{R^o} \int_{\underline{r}+\gamma_t}^{R^o} g_o(W_t-\gamma_t, R_t-\gamma_t)\,dW_t\,dR_t + \int_{R^o}^{W_{t-1}} \int_{\underline{r}+\gamma_t}^{R_t} g_o(W_t-\gamma_t, R_t-\gamma_t)\,dW_t\,dR_t$$

$$\equiv \ell_t^o(\gamma_t, R^o) + \ell_t^e(W_{t-1}, \gamma_t, R^o).$$

The layoff rate ℓ_t is the sum of the layoff rates to the nonmarket sector ℓ_t^o and to employment with the other firm ℓ_t^e. The total separation rate s_t--the sum of the quit and layoff rates--is the mass over the upper half-space plus the mass over a triangular region of the lower half-space.

$$(3.3) \qquad s_t(\gamma_t, R^o) = \int_{\underline{r}+\gamma_t}^{\bar{r}+\gamma_t} \int_{\underline{r}+\gamma_t}^{R_t} g_o(W_t-\gamma_t, R_t-\gamma_t)dW_t dR_t$$

$$+ \int_{\underline{r}+\gamma_t}^{R^o} \int_{R_t}^{R^o} g_o(W_t-\gamma_t, R_t-\gamma_t)dW_t dR_t.$$

The probability of continued employment is $1 - s_t$.

In this paper, I focus on the cyclical implications of the model.[4] Applying Leibniz's Rule for differentiating integral expressions, the following comparative statics result:

$$(4.1) \qquad \frac{\partial q_t}{\partial \gamma_t} = \int_{\underline{r}+\gamma_t}^{W_{t-1}} g_o(W_t-\gamma_t, W_{t-1}-\gamma_t)dW_t \equiv b(W_{t-1}, \gamma_t) > 0$$

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[4] The structural implications are analyzed in a more general context in McLaughlin (1987b). They are summarized as follows:

$$\frac{\partial q_t}{\partial W_{t-1}} < 0, \quad \frac{\partial \ell_t}{\partial W_{t-1}} > 0, \quad \frac{\partial \ell_t^o}{\partial W_{t-1}} = 0, \quad \frac{\partial \ell_t^e}{\partial W_{t-1}} > 0, \quad \frac{\partial s_t}{\partial W_{t-1}} = 0;$$

$$\frac{\partial q_t}{\partial R^o} = 0, \quad \frac{\partial \ell_t}{\partial R^o} > 0, \quad \frac{\partial \ell_t^o}{\partial R^o} > 0, \quad \frac{\partial \ell_t^e}{\partial R^o} < 0, \quad \frac{\partial s_t}{\partial R^o} > 0.$$

Neither the total separation rate nor the layoff rate to the nonmarket sector is influenced by the pre-separation wage W_{t-1}. But workers with higher wages are less (more) likely to have a subsequent separation labeled a quit (layoff). Productivity in the nonmarket sector is not a determinant of the quit rate but the total separation and layoff rates are increasing in R^o. Also, layoffs to the nonmarket sector relative to total layoffs are increasing in R^o.

$$(4.2) \qquad \frac{\partial \ell_t}{\partial \gamma_t} = -\int_{\underline{r}+\gamma_t}^{W_{t-1}} g_o(W_t-\gamma_t,\ W_{t-1}-\gamma_t)dW_t - \int_{\underline{r}+\gamma_t}^{R^o} g_o(R^o-\gamma_t,\ R_t-\gamma_t)dR_t$$

$$\equiv -\,[b(W_{t-1},\ \gamma_t) + c(\gamma_t,\ R^o)] < 0$$

$$(4.3) \qquad \frac{\partial s_t}{\partial \gamma_t} = -\int_{\underline{r}+\gamma_t}^{R^o} g_o(R^o-\gamma_t,\ R_t-\gamma_t)dR_t \equiv -\,c(\gamma_t,\ R^o) < 0.$$

where $b(W_{t-1}, \gamma_t)$ is the probability of getting a draw at the boundary between the Q and L regions, and $c(W_{t-1}, \gamma_t, R^o)$ at the vertical boundary between the L and CE regions. (See Figure 2.) The contemporaneous comovement with the business cycle is positive for quits, and negative for both layoffs and total separations.

The comparative static results in equations (4.1) - (4.3) hold R^o and W_{t-1} constant. Holding R^o constant is not a problem if the worker's nonmarket productivity is invariant to the business cycle (and constant or smoothly varying in the empirical analysis of section II). However, the ceteris paribus assumption regarding W_{t-1} does present a problem since W_{t-1} varies over the business cycle; that is, W_{t-1} and γ_t are correlated if business cycles exhibit "persistence." Furthermore, using aggregate data of the sort employed in most of section II below, one cannot condition on W_{t-1}. The solution is to integrate out W_{t-1}. Integrating over W_{t-1} with respect to its density--that is, the density of accepted wages in period t-1, $f_o(W_{t-1}-\gamma_{t-1})$ --the quit and layoff rates are expressed as functions of γ_t and γ_{t-1}.

$$(5.1) \qquad q_t(\gamma_t, \gamma_{t-1}) = \int_{\underline{r}+\gamma_{t-1}}^{\bar{r}+\gamma_{t-1}} q_t(W_{t-1}, \gamma_t)\cdot f_o(W_{t-1}-\gamma_{t-1})dW_{t-1}$$

$$(5.2) \qquad \ell_t(\gamma_t, \gamma_{t-1}, R^o) = \int_{\underline{r}+\gamma_{t-1}}^{\bar{r}+\gamma_{t-1}} \ell_t(W_{t-1}, \gamma_t, R^o)\cdot f_o(W_{t-1}-\gamma_{t-1})dW_{t-1}.$$

(The total separation rate $s_t(\gamma_t, R^o)$ is unaltered.) This modification does not affect the contemporaneous cyclicality of labor turnover. Differentiating expressions (5.1) and (5.2) with respect to γ_t yields (4.1) and (4.2) averaged over W_{t-1}.

$$(6.1) \qquad \frac{\partial q_t}{\partial \gamma_t} = \int_{\underline{r}+\gamma_{t-1}}^{\bar{r}+\gamma_{t-1}} b(W_{t-1}, \gamma_t)\cdot f_o(W_{t-1}-\gamma_{t-1})dW_{t-1} > 0$$

$$(6.2) \qquad \frac{\partial \ell_t}{\partial \gamma_t} = -\int_{\underline{r}+\gamma_{t-1}}^{\bar{r}+\gamma_{t-1}} b(W_{t-1}, \gamma_t)\cdot f_o(W_{t-1}-\gamma_{t-1})dW_{t-1} - c(\gamma_t, R^o) < 0.$$

The quit rate is decreasing in, and the layoff rate is increasing in γ_{t-1}. This follows since the pre-separation wage W_{t-1} is on average increasing in γ_{t-1}, and the quit (layoff) rate is decreasing (increasing) in W_{t-1}. Differentiating equations (5.1) and (5.2) with respect to γ_{t-1}, and integrating the resulting expressions by parts yields[5]

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[5] In writing equations (6.3) and (6.4), I use the following result:

$$(6.3)\qquad \frac{\partial q_t}{\partial \gamma_{t-1}} = - \int_{\underline{r}+\gamma_t}^{\bar{r}+\gamma_t} b(W_{t-1}, \gamma_t)\cdot f_o(W_t-\gamma_{t-1})dW_{t-1} < 0$$

$$(6.4)\qquad \frac{\partial \ell_t}{\partial \gamma_{t-1}} = \int_{\underline{r}+\gamma_t}^{\bar{r}+\gamma_t} b(W_{t-1}, \gamma_t)\cdot f_o(R_t-\gamma_{t-1})dW_{t-1} > 0.$$

Pulling these results together, I conclude: Quits and layoffs depend on both current and once-lagged states of the economy with opposite and serially switching signs. Higher-order lags are not determinants of the turnover variables. Regarding the quit rate, the effects of γ_t and γ_{t-1} enter as a difference; that is, only the business cycle shock ξ_t effects quits. This restriction does not apply to layoffs. The asymmetry results from some layoffs separating to the nonmarket sector. Total separations are countercyclical, but are independent of lagged cyclical fluctuations. These six comparative static results, and the restrictions they imply, are the principal implications of the model.

Additional Implications

The model generates several additional implications: First, the employment rate is procyclical. Second, only real business cycle fluctuations matter. Third, a distinction between anticipated and unanticipated cyclical shocks is irrelevant.

$$\frac{\partial q_t}{\partial W_{t-1}} = - b(W_{t-1}, \gamma_t).$$

The analysis to this point has been limited to labor force transitions (flows) without regard to participation rates in the various labor force states (stocks). Turn now to the participation rate in the market sector: the employment rate. To generate a prediction governing the cyclicality of employment, transitions into the market sector from the household must be modeled. Here I assume the separation rate from the household to the market sector is given by s_t^h, a function increasing in γ_t and decreasing in R^o. Employing the Markov assumption, the employment rate e_t is

$$(7.1) \qquad e_t = [1-\ell_t^o(\gamma_t, R^o)]e_{t-1} + s_t^h(\gamma_t, R^o)(1-e_{t-1}).$$

If the stochastic opportunities are independent of employment status, then $s_t^h = 1 - \ell_t^o$. Therefore,

$$(7.2) \qquad e_t(\gamma_t, R^o) = 1 - \ell_t^o(\gamma_t, R^o).$$

The employment rate is procyclical:

$$(8) \qquad \frac{\partial e_t}{\partial \gamma_t} = -\frac{\partial \ell_t^o}{\partial \gamma_t} = \int_{\underline{r}+\gamma_t}^{R^o} g_o(R^o-\gamma_t, R_t-\gamma_t)dR_t + \int_{\underline{r}+\gamma_t}^{R^o} g_o(W_t-\gamma_t, R^o-\gamma_t)dW_t > 0,$$

and is independent of the lagged employment rate e_{t-1} and the lagged cyclical state of the economy γ_{t-1}. That employment is procyclical is well known (see, e.g., Prescott et al. 1983, Table 2; Coleman 1984, Chapter 3), and is documented in section II below. Furthermore, the lag-structure implication is testable.

I maintain the assumption that all variables including W_{t-1} are measured in real terms (or inflated to current prices). Thus decomposing the nominal (percentage) variation into output and price components, $\gamma_t^n \equiv \gamma_t + \gamma_t^p$, it follows that increases in γ_t^p translate g_0 but also shift W_{t-1} and R^0.[6] Turnover can be shown to be neutral with respect to translations of g_0 by a constant and additions of the same constant to W_{t-1} and R^0. Therefore only real business cycle fluctuations are predicted to matter. I term this the price-neutrality implication.

A distinction between anticipated and unanticipated cyclical shocks is irrelevant. If I were to use the expectation at time t-1 of W_t as the benchmark used to divide separations into quits and layoffs, a distinction between anticipated and unanticipated cyclical shocks would be generated. However, I show in other work that the pre-separation wage benchmark is required to account for many of the empirical regularities which characterize turnover behavior (McLaughlin 1987a). Therefore, fully anticipated cyclical shocks are equivalent to unanticipated shocks in generating the cyclical movements of turnover.

In the following section, I document the cyclical comovements of labor turnover and employment with aggregate output, and in addition test the

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[6]That W_{t-1} and R^0 shift in response to cyclical price variation but remain fixed with respect to cyclical output variation requires justification. I assume that an optimal contract between a firm and a worker can automatically update the values of all variables in response to price shocks. However, this is not the case for cyclical output shocks. First, R^0 is invariant to cyclical output shocks. Second, initial asymmetries in the information structure preclude such automatic updating of W_{t-1} in response to cyclical output shocks. More precisely, if information common to the worker-firm match is not sufficient to disentangle idiosyncratic from aggregate shocks, there can be no automatic updating. See McLaughlin (1987b) for a treatment of how, in the context of the current model, initial informational asymmetries are resolved in a wage revision process.

countercyclical-separations, lag-structure, and price-neutrality implications. The other prediction awaits empirical testing elsewhere.

II. EMPIRICAL ANALYSIS

In this section, I use quarterly data for the post-war period and annual data from 1930 forward to document the cyclical behavior of aggregate labor turnover and employment rates in the U.S., and to test three additional implications of my model. I also aggregate micro-level panel data from the Panel Study of Income Dynamics (PSID) to control directly for the effect of the pre-separation wage on turnover. The section begins with a discussion of the aggregate data and methodology; the results for quarterly and annual analyses follow; I close this section with results from the PSID for the years 1968 - 1980.

Monthly data from the turnover survey of manufacturing establishments are published by the BLS for the period 1930-81. A respondent to the survey is asked to decompose total separations from the establishment into the following four categories: quits, discharges, layoffs, and other separations. (Published data include discharges in other separations.[7]) To guide the classification, the four categories are defined in the survey's instructions. For instance, quits are defined on the survey form as:

> ... a termination of employment initiated by the employee for any reason except to retire, to transfer

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[7] Unpublished monthly discharge data obtained from the BLS reveal that on average half of "other separations" are discharges, and the discharge rate is one-third the magnitude of the layoff rate. (The sample period is 1972:01 - 1981:12.) A startling finding is that the correlation between layoffs and discharges is -.52. This oddity remains even under the more formal empirical methodology employed in the text. In preliminary work, I estimate a cyclical discharge elasticity of 5.3: discharges are procyclical. This result is an anomaly since theoretical work suggests layoffs and discharges are similar. In my theory, they are identical.

> to another establishment of the same firm, or for service in the Armed Forces. (BLS 1971, 39)

The essential distinction between the definitions is the "initiated by" clause. Quits are initiated by the worker, and discharges and layoffs by the firm.

The turnover data are limited to the manufacturing sector. The amplitude of the cyclical deviations of manufacturing output exceeds that of GNP, that is, the manufacturing sector is overly procyclical. In my model, this corresponds to a translation of g_0 along a ray flatter than the 45° line (with manufacturing as the incumbent sector). The density moves toward the "continued employment" region. The countercyclicality of layoffs and separations is augmented, but the effect on quits is ambiguous. Although this potential bias is recognized, it is not taken into account in the estimation below.[8]

The empirical methodology is as follows. The aggregate turnover, employment, output, and price series are decomposed into secular and cyclical (i.e., smoothly varying and residual) components. A time-trend model predicts the secular movements. The residuals from these regressions are the variables of primary interest.[9] The residual components of the turnover and employment

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[8] Since the turnover data are for the manufacturing sector, one might conclude that the empirical counterpart to the cyclical productivity variation γ_t should be a manufacturing-specific cyclical variable. However, the theory applies to any individual, demographic group, occupation, or industry, as well as to the labor market as a whole; hence the cyclical component of aggregate (real) GNP is appropriate for the analysis of labor turnover in manufacturing.

[9] The analysis is limited to cyclical movements because the value of nonmarket production R^0 is likely to vary secularly. In addition, smoothly varying changes in the workforce, including its demographic composition, are factored out by detrending. Of course, one need not detrend the data series: the same basic estimates would result from regressions on "trended" series, but which include variables capturing time trends in the regressions.

series are regressed on the cyclical component(s) of the output (and price) series. The model is tested in three ways. The residual component of separations is expected to be negatively related to the business cycle measure. Lagged cyclical output fluctuations are introduced to test the lag-structure implication. To evaluate the price neutrality implication, I test for a zero cyclical price effect.

Since the theory does not guide the choice of a time frame, this methodology is applied to both quarterly and annual data. The results do document the strong cyclical movements of quits, layoffs, and employment. The test results are mixed, with the layoff rate regressions offering the most support for the joint wealth maximizing hypothesis.

Results from Quarterly Data

The secular components are estimated by a cubic function of time augmented for seasonal movements:[10]

$$\text{(9)} \qquad \log X_t = \alpha_0^x + \alpha_1^x\, t + \alpha_2^x\, t^2 + \alpha_3^x\, t^3 + \text{seasonals} + \epsilon_t^x,$$

where X = (T, Y, y, P) is a septuple of turnover and employment rate (T = (q, ℓ, s, e)), nominal (Y) and real (y) output, and price (P) series. Each of the seven regressions is estimated with an AR(1) correction for serially correlated residuals; the results are reported in Table 1. The secular or

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[10]The raw turnover and employment rate data are not-seasonally-adjusted monthly observations. Each series' quarterly average is used in the analysis in this section. The not-seasonally-adjusted quarterly GNP and the quarterly GNP price deflator are the nominal output and price variables.

Since the monthly civilian employment rate data (which are unpublished but obtainable from the BLS) are available from January 1948 forward, and the turnover survey was terminated at the close of 1981, the sample time frame is 1948:I - 1981:IV.

TABLE 1

ESTIMATES OF SECULAR TRENDS
--Quarterly Data--

	Dependent Variable	Constant	t	t^2	$t^3/100$	Seasonals	$\hat{\rho}$	R^2
(1)	QUITS	1.15 (0.23)	−0.206 (0.059)	0.014 (0.004)	−0.026 (0.008)	yes [235.0]	.89 (.04)	.57
(2)	LAYOFFS	0.33 (0.21)	0.089 (0.054)	−0.007 (0.004)	0.014 (0.007)	yes [17.8]	.70 (.06)	.31
(3)	SEPARATIONS	1.62 (0.06)	−0.060 (0.016)	0.004 (0.001)	−0.007 (0.002)	yes [61.8]	.53 (.07)	.57
(4)	EMPLOYMENT	4.01 (0.01)	−0.001 (0.004)	−0.000 (0.000)	0.000 (0.000)	yes [210.8]	.87 (.04)	.64
(5)	GNP	7.79 (0.03)	0.061 (0.007)	−0.001 (0.000)	0.003 (0.001)	yes [204.4]	.80 (.05)	.998
(6)	P	−0.66 (0.02)	0.027 (0.004)	−0.001 (0.000)	0.004 (0.001)	yes [0.6]	.91 (.04)	.998
(7)	GNP − P	8.45 (0.03)	0.033 (0.007)	0.000 (0.000)	−0.001 (0.001)	yes [233.0]	.82 (.05)	.992

Note: 136 observations per regression. The dependent variables are measured in logs. The standard errors are in parentheses. The F-statistics corresponding to the joint tests of the seasonal dummies are in brackets. At the 5% level, the critical value for F(3, 128) is 2.68.

predicted series are nonmonotonic for the turnover variables. The quit and separation rates behave similarly, reaching a trough in 1959 and peaking in 1973. Layoffs peak in 1956 and reach a trough in 1975. The employment rate is flat until 1960, and then rises monotonically. The secular output and price series are monotonically increasing. The real output variable grows at a falling rate, but the nominal series grows at an increasing rate due to rising inflation.

My primary interest is in the properties of the residuals of these cubic regressions: $\{\hat{\epsilon}_t^x\}$. The residual turnover and employment series $\{\hat{\epsilon}_t^x\}$ are plotted with the cyclical component of real GNP $\{\hat{\epsilon}_t^y\}$ in Figures 3-6. The procyclical nature of both quits and employment, and the countercyclical nature of layoffs are quite visible. Inspection of Figure 5 reveals no obvious cyclical pattern for separations.

In Table 2, I report the results of AR(1) regressions of the residual turnover and employment series on the business cycle variable.

$$(10) \qquad \hat{\epsilon}_t^T = \beta_0^T + \beta_1^T \hat{\epsilon}_t^y + u_t^T; \qquad\qquad u_t^T = \rho^T u_{t-1}^T + \nu_t^T.$$

The regression results corroborate the graphical evidence. Regressions (1), (4), (7), and (10) indicate that quits and employment are procyclical, layoffs countercyclical, and separations acyclical. The quit regression implies that a business cycle expansion of 1 percent results in a contemporaneous rise in the quit rate of 5.2 percent. The layoff, separation, and employment rate regressions yield elasticities of -7.3, -0.05, and 0.2, respectively.[11] The

[11]Under constant returns to scale and with other factors of production fixed, the employment elasticity is the inverse of "employments's share": the estimated coefficient in regression (10) should exceed one. However, see Lucas (1970) for one way (varying capital utilization) to reconcile this well-known result with producer theory.

Figure 3

PROCYCLICAL QUIT RATE

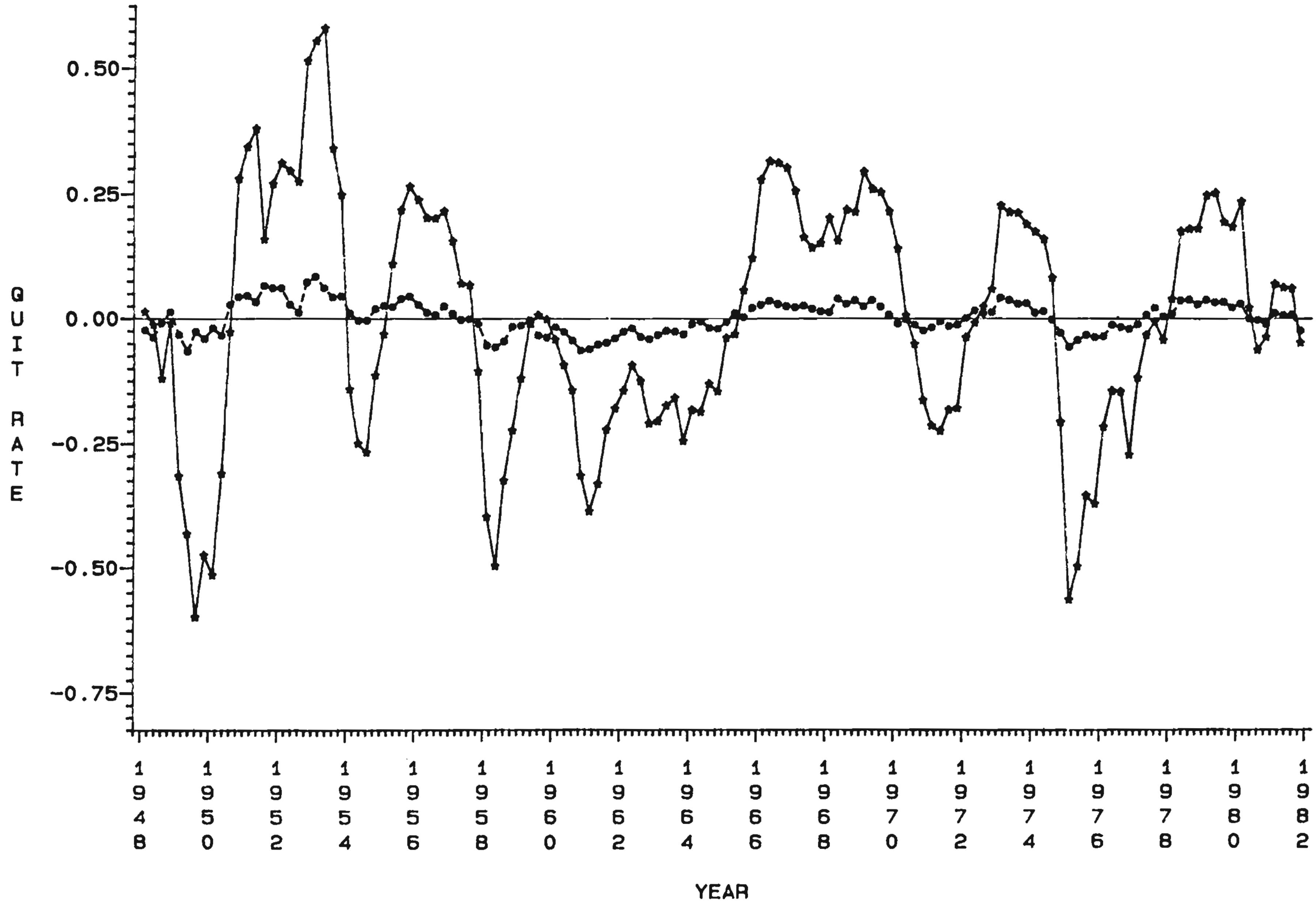

Figure 4

COUNTERCYCLICAL LAYOFF RATE

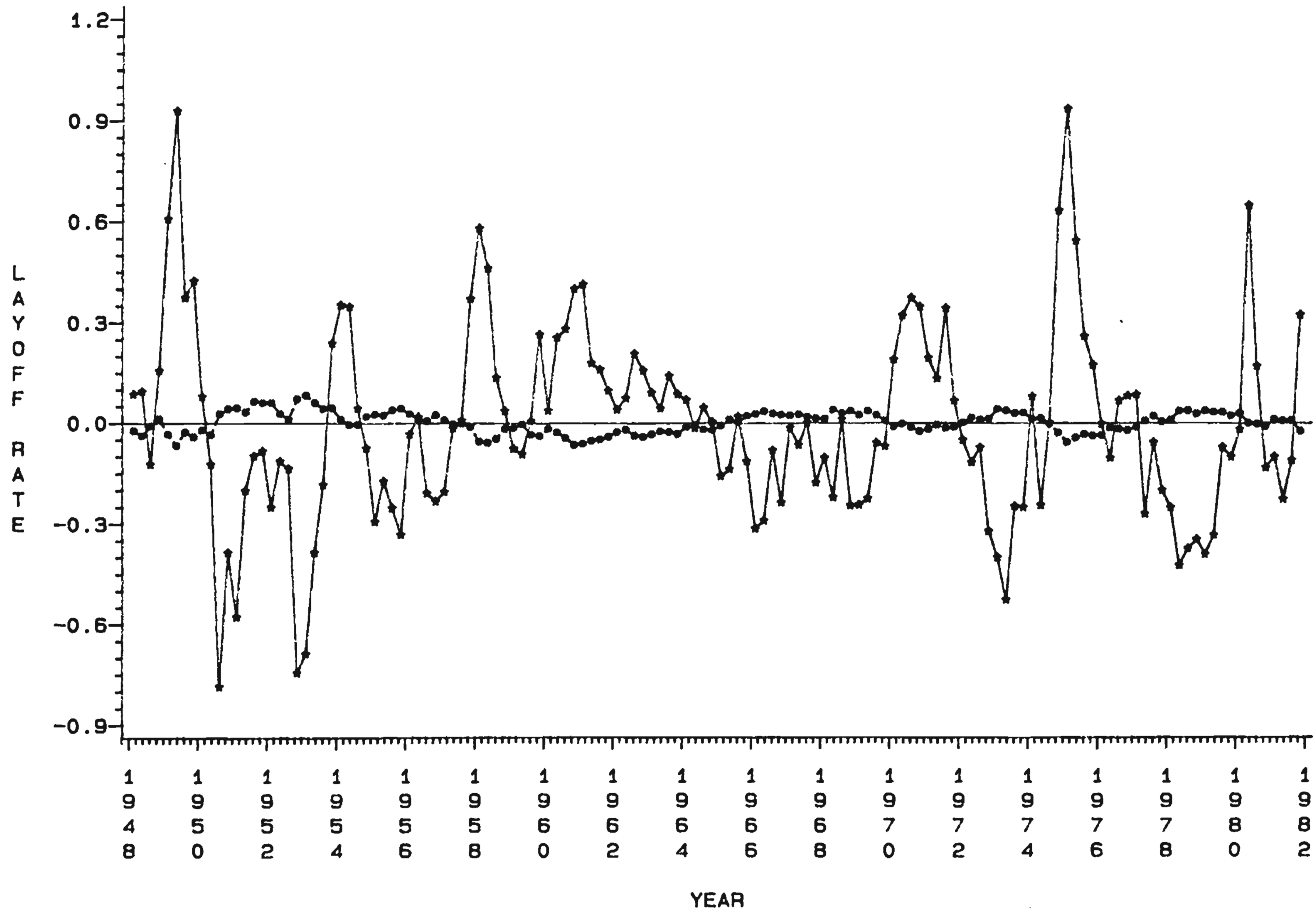

Figure 5

TOTAL SEPARATION RATE

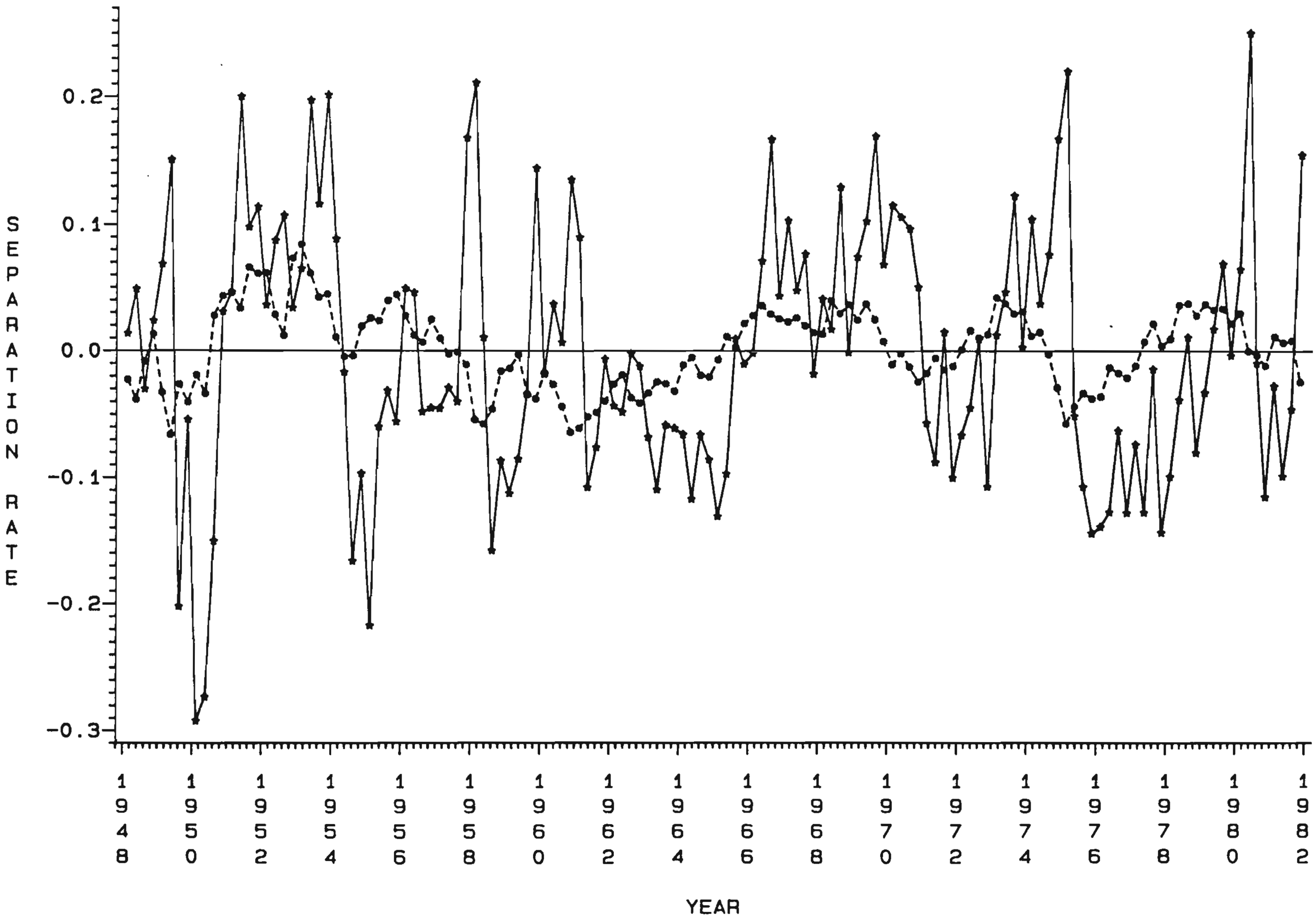

Figure 6

PROCYCLICAL EMPLOYMENT RATE

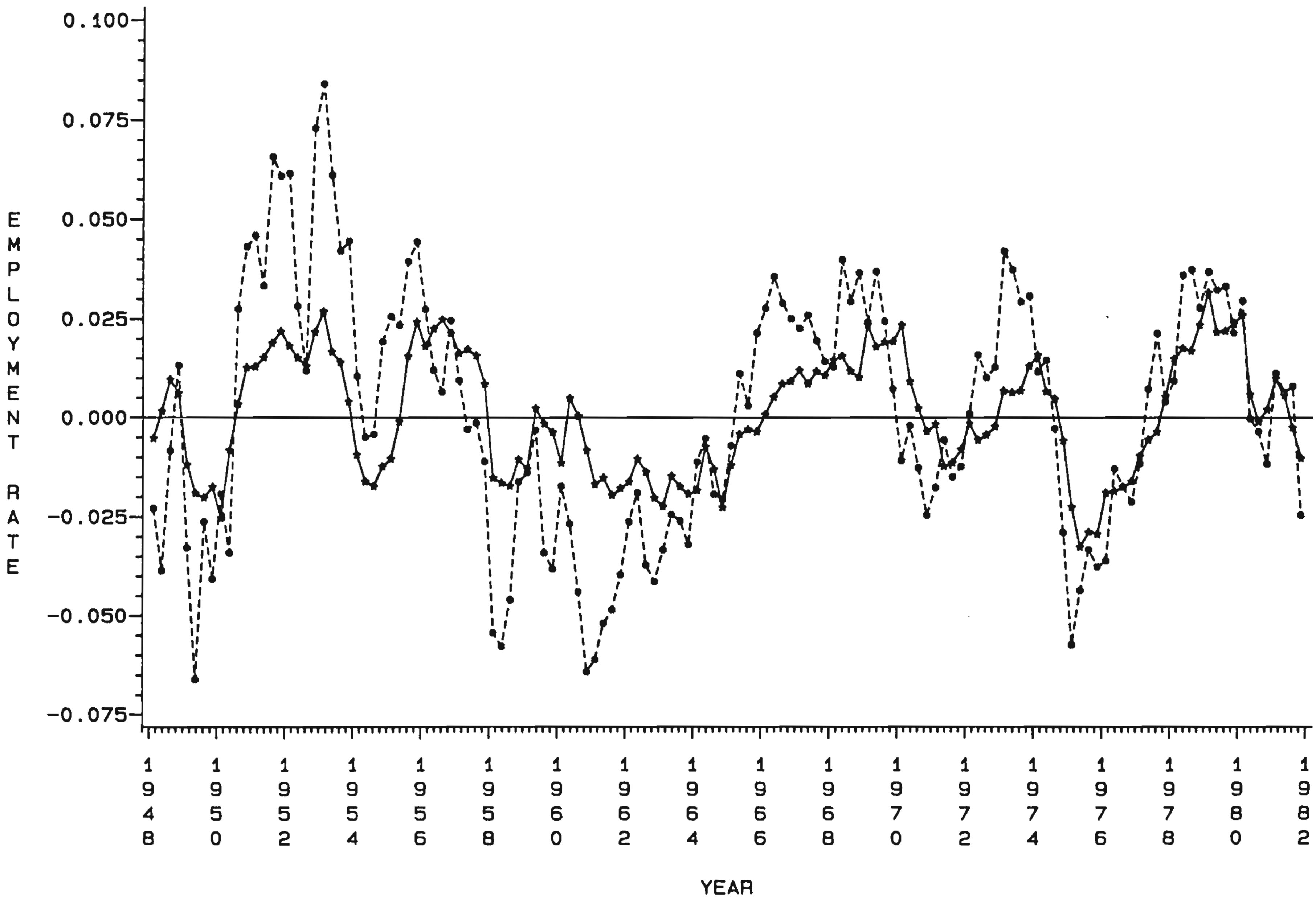

TABLE 2

ESTIMATES OF TURNOVER CYCLICALITY
--Quarterly Data--

	Dependent Variable	N	$RGNP_t$	$RGNP_{t-1}$	P_t	P_{t-1}	$\hat{\rho}$	R^2	Test Statistics
(1)	QUITS	136	5.17				.55	.73	
			(0.41)				(.07)		
(2)	QUITS	135	3.77	2.76			.65	.80	125.4 [D]
			(0.40)	(0.40)			(.07)		
(3)	QUITS	135	3.87	2.82	1.19	-0.21	.63	.81	0.6 [PN]
			(0.40)	(0.40)	(1.00)	(1.00)	(.07)		42.0 [D, PN]
(4)	LAYOFFS	136	-7.29				.56	.58	
			(0.70)				(.07)		
(5)	LAYOFFS	135	-8.77	2.64			.41	.63	39.0 [D]
			(0.77)	(0.78)			(.05)		
(6)	LAYOFFS	135	-8.83	2.90	-4.29	5.28	.41	.65	3.6 [PN]
			(0.77)	(0.77)	(1.94)	(1.96)	(.08)		15.9 [D, PN]
(7)	SEPARATIONS	136	-0.05				.48	-.01	
			(0.36)				(.08)		
(8)	SEPARATIONS	135	-1.47	2.58			.37	.33	
			(0.39)	(0.39)			(.08)		
(9)	SEPARATIONS	135	-1.39	2.66	0.38	0.62	.36	.35	1.5 [PN]
			(0.39)	(0.39)	(0.99)	(1.00)	(.08)		13.5 [C, PN]
(10)	EMPLOYMENT	136	0.23				.72	.50	
			(0.03)				(.06)		
(11)	EMPLOYMENT	135	0.20	0.13			.73	.61	
			(0.03)	(0.03)			(.06)		
(12)	EMPLOYMENT	135	0.21	0.13	0.01	-0.01	.71	.63	0.4 [PN]
			(0.03)	(0.03)	(0.08)	(0.08)	(.06)		6.3 [C, PN]

NOTES TO TABLES 2 AND 3

All variables are in logs and are detrended. Standard errors are in parentheses. The final columns of the two tables report the F-statistics corresponding to tests of various restrictions. In brackets are labels identifying the restrictions: [D] differences; [PN] price neutrality; and [C] contemporaneous (no lagged effect). The critical values for the F-tests at the 5 percent level are:

$F(1, 131) = 3.92$	$F(1, 47) = 4.00$
$F(2, 129) = 3.07$	$F(2, 46) = 3.15$
$F(3, 128) = 2.68$	$F(3, 46) = 2.76$.

acyclicality of the separation rate conflicts with the countercyclical separations implication.

I employ variations on equation (10) to test the lag-structure and price-neutrality implications.

$$(11) \qquad \hat{\epsilon}^{T}_{t} = \beta^{T}_{0} + \beta^{T}_{1}\,\hat{\epsilon}^{y}_{t} + \beta^{T}_{2}\,\hat{\epsilon}^{y}_{t-1} + \beta^{T}_{3}\,\hat{\epsilon}^{P}_{t} + \beta^{T}_{4}\,\hat{\epsilon}^{P}_{t-1} + u_{t}.$$

Based on the theoretical cyclicality results in equations (6.1) - (6.4), (4.3) and (8), the predicted sign configuration in regressions (2), (5), (8), and (11) is:

$$(12.1) \qquad \beta^{q}_{1} > 0, \qquad \beta^{q}_{2} < 0, \qquad \text{with} \qquad \beta^{q}_{1} + \beta^{q}_{2} = 0;$$

$$(12.2) \qquad \beta^{\ell}_{1} < 0, \qquad \beta^{\ell}_{2} > 0, \qquad \text{with} \qquad \beta^{\ell}_{1} + \beta^{\ell}_{2} < 0;$$

$$(12.3) \qquad \beta^{s}_{1} < 0, \qquad \beta^{s}_{2} = 0;$$

$$(12.4) \qquad \beta^{e}_{1} > 0, \qquad \beta^{e}_{2} = 0.$$

The layoff rate regression offers support for the lag-structure implication: the contemporaneous cyclical effect is negative and the lagged cyclical effect is positive and smaller in magnitude. The lag structure implication does not find empirical support from the other three regressions. The cyclical effects do not switch signs serially in the quit rate regression. At odds with predictions (12.3) and (12.4), the lagged output variable is significant in both separation and employment rate regressions. However, the predicted countercyclicality of total separations is observed, as predicted, in regression (8).

The elasticities associated with cyclical price fluctuation are substantially smaller than those associated with cyclical output fluctuation. The price-neutrality implication is evaluated formally by testing the restriction of $\beta_3^T = \beta_4^T = 0$ for quits, layoffs, separations, and employment.[12] The results are reported as regressions (3), (6), (9), and (12) in Table 2. In only the layoff rate regression is the null hypothesis rejected, and there only marginally. For quits, separations, and employment, individual t-tests of the significance of the price variables fail to reject the null of the price neutrality implication. In addition, F-tests for each regression also fail to reject the price neutrality null. However, for layoffs the elasticities are sufficiently large to reject price neutrality even though the contemporaneous effect of cyclical price fluctuation is less than one-half the magnitude of the contemporaneous effect of cyclical output fluctuation.

The empirical results from quarterly data reveal the expected strong cyclical movements in quits, layoffs, and employment. The evidence also indicates weak countercyclical variation in separations. While the results on the lag structure implications are mixed at best, the evidence supports the price neutrality implication.

Results from Annual Data

The same empirical methodology is applied to the annual data for the

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[12] One can also test the hypothesis of a nominal business cycle variation; that is, $\beta_1^T = \beta_3^T$ and $\beta_2^T = \beta_4^T$. In Tables 2 and 3, this is rejected in nearly every case.

period 1930 to 1981, but I omit the results of the detrending procedure. The results of AR(1) cyclical regressions are summarized in Table 3. The cyclical quit and layoff elasticities are somewhat smaller than their corresponding quarterly magnitudes. The elasticities of the contemporaneous cyclical effects are 3.5 for quits, -2.1 for layoffs, 0.4 for separations, and 0.5 for employment. In contrast with the results from quarterly data, the separation rate is significantly procyclical.

The lagged cyclical effect is, as predicted, negative for quits and positive for layoffs; however, the effect is not significant for quits and is only borderline significant for layoffs. As with the quarterly data, the annual results clearly reject the lag structure implication for quits: $\beta_1^q + \beta_2^q \neq 0$; however, as predicted, $\beta_1^\ell + \beta_2^\ell < 0$. That the lagged real output variable is statistically insignificant in the separation and employment rate regressions also is as predicted. Consequently, regressions (5), (8), and (11) offer clear support for this implication of the theory.

Tests of the price-neutrality implication on the annual data differ considerably from the test results on quarterly data. Although both layoffs and separations behave as predicted, exhibiting a zero price response, strong price effects are exhibited in quit rate regression (3) and employment rate regression (12).[13]

As with the results form quarterly data, the estimates of contemporaneous cyclical effects on annual data are quite strong for quit, layoff, and employment rates. Although the countercyclical separation implication does

[13] The serial correlation of the price variable is substantial (0.94 in quarterly data and 0.93 in the annual data) indicating a collinearity problem. In every case where price neutrality is rejected by the joint test, it is not rejected when either, rather than each, of the two price variables is included in the regressions.

TABLE 3

ESTIMATES OF TURNOVER CYCLICALITY
--Annual Data--

	Dependent Variable	N	$RGNP_t$	$RGNP_{t-1}$	P_t	P_{t-1}	$\hat{\rho}$	R^2	Test Statistics
(1)	QUITS	52	3.51				.67	.66	
			(0.47)				(.11)		
(2)	QUITS	51	3.69	-0.24			.61	.65	26.7 [D]
			(0.62)	(0.59)			(.12)		
(3)	QUITS	51	3.36	0.26	3.00	-3.43	.57	.77	9.4 [PN]
			(0.54)	(0.52)	(0.80)	(0.78)	(.12)		18.4 [D, PN]
(4)	LAYOFFS	52	-2.05				.18	.36	
			(0.45)				(.14)		
(5)	LAYOFFS	51	-3.14	1.32			.20	.40	9.5 [D]
			(0.72)	(0.70)			(.14)		
(6)	LAYOFFS	51	-2.91	1.03	-0.41	0.88	.17	.43	0.8 [PN]
			(0.75)	(0.72)	(1.07)	(1.06)	(.15)		3.8 [D, PN]
(7)	SEPARATIONS	52	0.39				.21	.10	
			(0.17)				(.14)		
(8)	SEPARATIONS	51	0.05	0.37			.17	.16	
			(0.26)	(0.25)			(.14)		
(9)	SEPARATIONS	51	-0.06	0.41	0.58	-0.49	.11	.21	1.1 [PN]
			(0.27)	(0.26)	(0.38)	(0.38)	(.15)		1.5 [C, PN]
(10)	EMPLOYMENT	52	0.47				.76	.73	
			(0.04)				(.09)		
(11)	EMPLOYMENT	51	0.45	-0.02			.80	.74	
			(0.04)	(0.04)			(.09)		
(12)	EMPLOYMENT	51	0.41	-0.02	0.20	0.00	.50	.95	28.0 [PN]
			(0.03)	(0.03)	(0.04)	(0.04)	(.13)		26.9 [C, PN]

not fare well in the annual data, there is evidence to support the lag structure implication. The price neutrality implication finds stronger support on the quarterly data.

Results from the PSID

A third source of evidence is aggregated data from the Panel Study of Income Dynamics (PSID). An advantage of these data is that for each individual in the sample the pre-separation wage is known. Thus in the aggregate one can condition on the average pre-separation wage rather than relying on the lagged cyclical fluctuation. However, with these data, one can be more precise. Under my hypothesis the fraction of separations who take higher paying jobs--the rate of positive wage growth f--is an exact predictor of the fraction of separations labeled quits--the conditional quit rate $\bar{q}$ (McLaughlin 1987a).

In Figure 7, the conditional quit rate, rate of positive wage growth, and business cycle time series are depicted.[14] One conclusion is immediate: both the conditional quit rate and the rate of positive wage growth are procyclical. (The correlations with the cyclical variable are, respectively, .92 and .87.) Consequently, the rate of positive wage growth is a likely candidate to explain turnover cyclicality. Regression (1) of Table 4 documents the cyclicality of the conditional quit rate in the aggregated PSID. In regression (2), I replace the cyclical variable with the rate of positive wage growth; this variable alone accounts for three-fourths of the squared variation in the conditional quit rate. Including both the cyclical and the rate of positive wage growth variables in regression (3), only the cyclical

[14] Construction of the aggregate time series from the PSID is described in McLaughlin (1987a). Note that the two series from the PSID are not detrended.

Figure 7

CONDITIONAL QUITS AND POSITIVE WAGE GROWTH

1968 - 1980

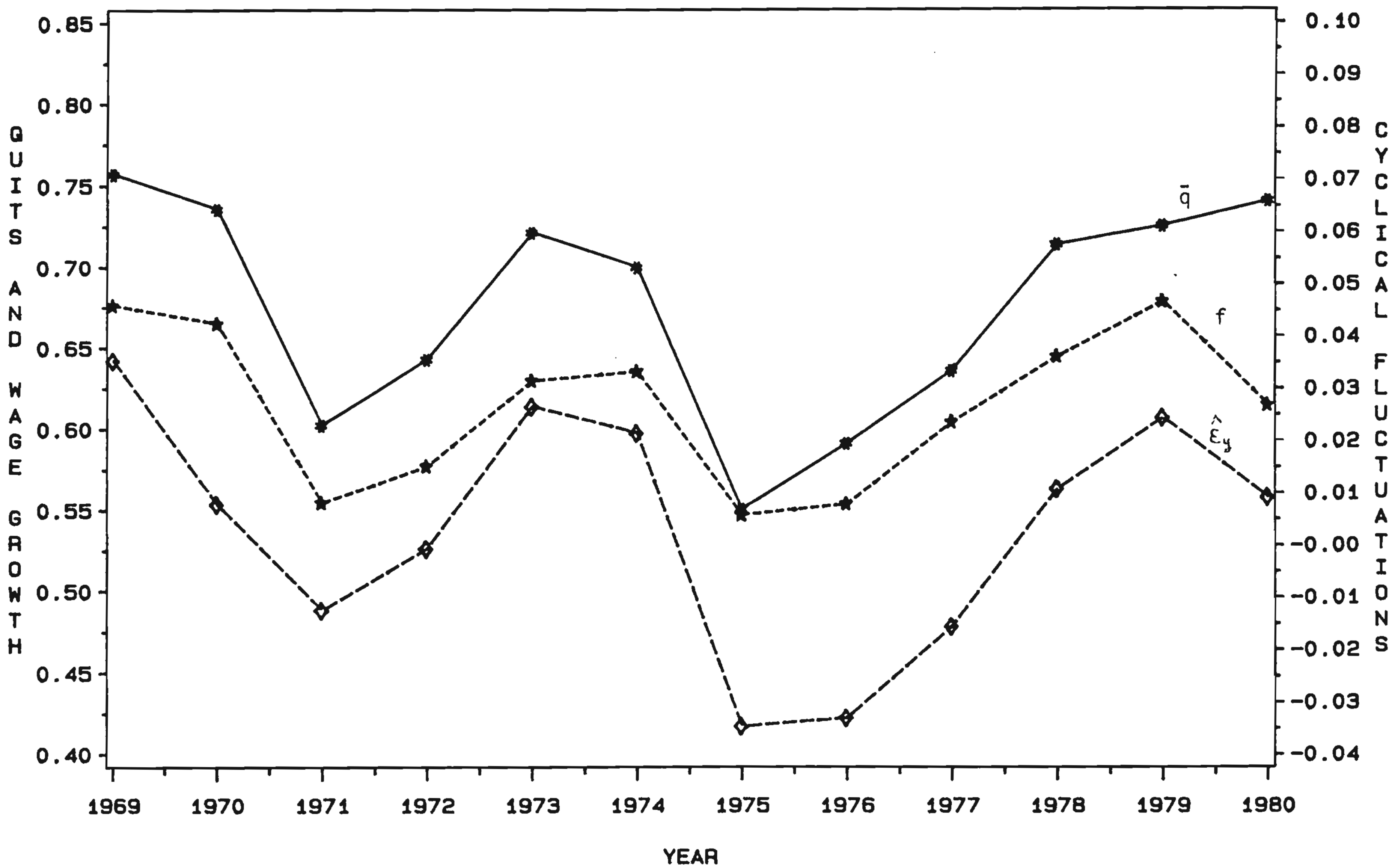

TABLE 4

ESTIMATES OF THE CONDITIONAL QUIT RATE
--Aggregated PSID Data--

	Dependent Variable	constant	$RGNP_t$	f_t	R^2	DW
(1)	CONDITIONAL QUITS	0.67 (0.01)	2.71 (0.40)		.82	1.72
(2)	CONDITIONAL QUITS	-0.11 (0.13)		1.27 (0.22)	.78	1.77
(3)	CONDITIONAL QUITS	0.33 (0.22)	1.71 (0.75)	0.55 (0.36)	.86	2.03

Note: The regressions are estimated as linear probability models using the minimum χ^2 estimator. There are 12 annual observations per regression from 3,310 underlying observations. Asymptotic standard errors are in parentheses.

variable remains as a statistically significant determinant of $\bar{q}$. This suggests that the rate of positive wage growth is not the only factor which varies over the business cycle and affects the conditional quit rate. Nevertheless, time series variation in the conditional quit rate is closely approximated by time series variation in the fraction of separations moving to higher paying jobs.

Taking Stock

The empirical analysis presented in this section indicates that my specification of the joint wealth maximizing hypothesis does reasonably well in capturing the principal features of turnover and employment cyclicality: quits and employment are procyclical, layoffs countercyclical, and the conditional quit rate moves roughly one-for-one with the rate of positive wage growth; tests of countercyclical separations, the lag structure, and price neutrality are mixed. Even where the evidence is counter to the predictions of my hypothesis, the results are instructive: the sign configurations across equations are quite rich. The results displayed in Tables 2 and 3 present a challenge to the researcher of turnover cyclicalilty

Consider as an alternative a rigid wage model, in particular Hall and Lazear's (1984) fixed-wage contract. The principal difference between the model presented in section I above and Hall and Lazear's fixed wage contract is that in the latter there are inefficient separations. But this alone has little impact on the cyclical implications. (See Hall and Lazear for definitions of the turnover regions under such a contract.) If their model is augmented for a nonmarket sector, a business cycle translation would generate countercyclical separations. (Without the nonmarket sector, it would fail to predict procyclical employment.) Turning to the lag structure, the model with a fixed wage fails to generate any effect for lagged cyclical fluctuations (once conditioned on the contemporaneous state of the economy). A partial-adjustment process for the wage would generate lagged effects, but is not likely to match up well against the observed diversity across the turnover and employment variables. Finally, a contract fixing a wage in nominal terms is rejected by the data (see note 12), but a fixed real-wage contract would yield the price neutrality implication. On balance, a simple fixed wage model is not consistent with the evidence in Tables 2 and 3.

Even if a rigid wage model could account for the cyclical regularities, it would remain deficient theoretically. Would the wage in the optimal contract be rigid with respect to business cycle shocks? The answer is yes only to the extent that information regarding variation in γ_t is asymmetric ex post. Hall and Lazear (1984) and Hashimoto and Yu (1980) contend that fully flexible wage contracts are suboptimal in the presence of asymmetric information. At the individual level informational asymmetries due to idiosyncratic shocks are likely to be important; nevertheless, is it likely that variation in γ_t is either (i) forecastable ex ante, (ii) verifiable ex post, or (iii) well proxied by contemporaneous observables such as the unemployment rate? If these elements are important in limiting the degree of ex post informational asymmetry between firms and workers, the wage in the optimal contract is largely flexible (Hashimoto and Yu 1980).[15] Without wage rigidity over the business cycle, this alternative model does not capture the cyclical differences between quits and layoffs.[16]

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[15]Of course, with risk averse workers and risk neutral firms variation in the wage would be limited as a form of insurance. Nevertheless, with symmetric information, the optimal contract would condition the separation decision on γ_t.

[16]In McLaughlin (1987a), I develop a similar argument in the life cycle

III. SUMMARY AND CONCLUSIONS

Treating quits and layoffs as inconsequential labels which are applied based on the pre-separation division of the match value is potentially fruitful in characterizing the cyclical properties of labor turnover. Quits are worker-initiated separations resulting in wage improvements, and layoffs are firm-initiated separations resulting in wage cuts; but both forms of separation are always joint wealth maximizing. Cyclical fluctuations are modeled as a translation of the joint density of the worker's productivity on his current job and his productivity in other firms. The probability of getting a draw in the quit (layoff) region, based on the pre-separation division of the match value, is shown to be increasing (decreasing) in the level of the business cycle.

The empirical results in section II document the implied cyclical regularities of quits, layoffs, and employment. In addition, using data aggregated from the Panel Study of Income Dynamics, I show that the quit rate conditional on separating moves close to one-for-one with the fraction of workers separating to higher paying jobs; and this "rate of positive wage growth" variable alone captures most of the cyclical variation in the conditional quit rate. On the more standard aggregate-turnover data, I find some evidence to support the countercyclical-separations and price-neutrality implications. Only the layoff rate supports the lag structure implication.

I have presented a simple model: neither the approach nor the specification are likely to be complete characterizations of turnover in the labor market. However, in testing the additional implications of the model, I

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context. There I show that the rigid wage <u>profile</u> generated in the optimal contract cannot account for the turnover regularities over the life cycle.

have gone beyond other models of cyclical turnover. To be successful, future research must confront these implications and empirical regularities in addition to the direct cyclical movements in quits and layoffs.

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