The Quit-Layoff Distinction: Growth Effects

McLaughlin, Kenneth J.

Working Paper No. 73
February 1987

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Kenneth J. McLaughlin

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In this paper, I account for several empirical findings regarding gains to labor mobility and the quit-layoff distinction. My analysis focuses on two issues: The first is differential wage growth by turnover type (quit or layoff): how this varies over the life cycle, with the steepness of the life-cycle wage profile, and with economy-wide productivity growth. The second issue is the effect of productivity growth on quit and layoff rates.

The regularities are:

(i) Wage gains to employment separations differ by the type of separation (quit or layoff): relative to stayers, quits exhibit higher, and layoffs lower wage growth in employment transitions (Bartel and Borjas 1981; Antel 1985; Mincer 1986; McLaughlin 1986a).

(ii) Wage growth is decreasing in experience (or age) and employment tenure, is increasing in education, is lower for union members, and covaries positively with cyclical "shocks" (McLaughlin 1986a; Bils 1985).

(iii) With the direct effects controlled for, turnover type <u>interacts</u> in the wage growth regression with experience, education, union status, and cyclical "shocks" (Mincer 1986; McLaughlin 1986a).

(iv) The ratio of quits to layoffs falls with experience, increases with education, and is higher for nonunion, and white workers (Leighton and Mincer 1982, 242; Mincer 1986; McLaughlin 1986a).

Explanations to account for several of these findings are not difficult to formulate; accounting for all the regularities is more difficult, but not intractable.

The paper is organized as follows: The analysis begins in section I with an overview of the wage growth regression and the correspondence between the listed empirical regularities and the parameters of this regression. The remainder of the paper includes my discussion of various interpretations and models to explain the results. These include a summary of Mincer's (1986) hypothesis in section II, and in section III a turnover model which admits various determinants of whether a given separation is a quit or layoff. I consider a purely random scheme and two turnover "labels" models. I find that the labels model employed in my other work on the quit-layoff distinction can account for items (i)-(iv).

My model incorporates the following features: Quits are separations to higher paying jobs, layoffs to lower paying jobs. General productivity growth increases (decreases) the probability of going to a higher (lower) paying job, hence quits are increasing relative to layoffs in general productivity growth. With a concave life-cycle productivity profile, productivity growth falls with experience or age, and consequently the quit-layoff ratio falls over the life cycle. The same parameterization implies that, relative to layoffs, quits are increasing in education and are relatively more frequent for white workers. The key is that the steepness of the life-cycle wage profile is increasing in education, and the profile is steeper for white workers. The model also applies to cyclical and secular variations.

The concave profile alone can account for wage gains to separations declining for both quits and layoffs over the life cycle; moreover, wage gains for both quits and layoffs are predicted to increase with education, to be higher for white workers, and to covary positively with cyclical shocks. However, the interaction effects are more complicated. A selectivity bias is introduced by growth as some relatively low wage-growth separations (would-be layoffs) are labeled quits. The bias is shown to interact with the respective quit and layoff rates in the manner required by item (iii).

The results are limited to the effect of general productivity growth. The model does not predict any effect of employment tenure on the ratio of quits to layoffs or through turnover-type interactions in wage growth regressions.

Also in section III, I discuss whether the more standard approach to the quit-layoff distinction, based on rigid wages, captures the growth-related regularities. Limiting the treatment to optimal contracting models of wage rigidity, I conclude that rigid wage models cannot account for the regularities: in the optimal contract, the wage is flexible with respect to growth in general productivity.

In section IV, I test a structural prediction of the model on a sample drawn from the Panel Study of Income Dynamics. While a formal test rejects the prediction of my model, the specification does reasonably well in accounting for time series variation in the fraction of separations labeled quits.

As a guide to empirical work on wage growth by turnover type, these results suggest the simultaneous equations specification described in the appendix. The specification can also be used to test my hypothesis by controlling for the endogeneity of the quit-layoff labels in the wage growth regression. Section V contains my conclusions.

I. REGRESSION FRAMEWORK

Consider the following wage growth equation to be estimated on a sample of both movers and stayers:

$$(1.0) \qquad \Delta W_i = \tilde{\gamma}_i + \sigma_Q Q_i + \sigma_L L_i + (Q_i G_i)\sigma_Q^G + (L_i G_i)\sigma_L^G , \qquad i = 1, \ldots, n,$$

where ΔW_i is the time difference of worker i's log wage, $\tilde{\gamma}_i$ is that component of productivity growth which is independent of turnover status, Q_i and L_i are dummy variables for quits and layoffs, and G_i is a set of interaction variables. Replacing $\tilde{\gamma}_i$ with its empirical counterpart, the regression $\tilde{\gamma}_i = X_i\beta + \epsilon_i$, equation (1.0) is rewritten[1]

$$(1.1) \qquad \Delta W_i = \alpha + X_i\beta + \delta_Q Q_i + \delta_L L_i + (Q_i G_i)\delta_Q^G + (L_i G_i)\delta_L^G + \epsilon_i \;, \qquad i = 1, \ldots, n.$$

X_i includes measures of economy-wide cyclical and secular growth; contingent on concave life-cycle productivity profiles, such variables as experience or age, and employment tenure are included in X_i; also included are variables that capture slope differences across productivity profiles, such as education, and indicators of union status, race and sex.

Evidence regarding the wage growth regularities are based on estimates from regressions of the general form of (1.1). First, omitting both the productivity growth regressors X_i and the interaction effects, estimates of α, δ_Q, and δ_L can be used to test the hypothesis of positive wage growth for quits and negative wage growth for layoffs: $\alpha + \delta_Q > 0$, and $\alpha + \delta_L < 0$. Including the direct effects, regularity (i) corresponds to $\delta_Q > 0$ and $\delta_L < 0$. Regularity (ii) is based on the signs of the various elements in β. With the interaction variables included, δ_Q^G and δ_L^G capture turnover-type interactions. For instance, consider the case of turnover-type - life-cycle interactions. Letting E denote labor market experience, $\delta_Q^E < 0$ and $\delta_L^E < 0$ implies that wage gains to both quits and layoffs decline over the life cycle. Differences between quits and layoffs are suppressed to make mover-stayer comparisons by imposing the restrictions $\delta_Q = \delta_L$ and $\delta_Q^G = \delta_L^G$ for all interaction variables.

II. MINCER'S HYPOTHESIS

Jacob Mincer's (1986) explanation of empirical regularities (i)-(iv), specifically the life-cycle relationships, turns on the differential incidence (and less importantly, differential duration) of unemployment by turnover type (quit or layoff). The foundation for the link between the two is search theory; specifically, unemployed search versus on-the-job search (OJS). Mincer associates quits with separations via OJS; layoffs, which presumedly occur exogenously, experience intervening spells of unemployment, thus a subsequent transition to employment is governed by unemployed search. A well-known theoretical proposition establishes that in a stationary environment the reservation wage of on-the-job searchers (quits) exceeds that of unemployed searchers (layoffs) (Burdett 1978). Therefore, the average accepted wage of quits exceeds that of layoffs. This accounts for item (i), the differential wage growth by turnover type.

The second step is to explain why wage gains from separating (changing employers) fall over the life cycle. Mincer offers one trivial explanation: the quit-layoff ratio falls with age, consequently the relative weight associated with low wage-growth transitions (layoffs) grows with age. This results in lower wage growth to the average separation. Mincer does not leave the systematic movement in the quit-layoff ratio unexplained, again resorting to OJS theory. If, as assumed by Mincer, the arrival rate of wage offers declines with age, the theory of OJS predicts longer spells of employment, meaning fewer quits.[2] There is no mechanism driving a change in the <u>incidence</u> of unemployment (layoffs). Hence quits fall, layoffs are constant, and the quit-layoff ratio falls with age.

In Mincer's discussion, there is an additional force reducing--with age--the average wage gain from separating. The reduced arrival rate of wage offers lowers the reservation wage of unemployed searchers, causing the average accepted wage of layoffs to fall. (There is a related, albeit ambiguous, effect on the duration of unemployment spells.) Consequently, the average wage gain of separations would fall with age even if the quit-layoff ratio were constant over the life cycle.

A problem with Mincer's arrival-rate hypothesis is that it does not account for the decline in wage gains to quits with age. Mincer confronts this final regularity by adopting a "push" versus "pull" approach to quits. That is, not all quits are "pull" (high wage growth) as is implicitly assumed in applying the OJS framework. Rather quits may include some "push" or "personal" (low wage growth) separations, and the fractions of these quit types may vary with age. Mincer concludes that young quits are primarily "pull," with the fraction of "push" and "personal" quits rising with age.

III. TURNOVER FRAMEWORK

The approach I take in accounting for these wage-growth - turnover-type regularities is based on a model of efficient turnover (see McLaughlin 1986b). 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 feature. 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 the initial period the representative worker is matched with a firm paying a (log) wage W^{o}, and this worker has an opportunity (log) wage with a second firm of R^{o}.[3] Between periods idiosyncratic shocks, which are common knowledge, arrive changing match values. These wage offers are drawn from a bivariate density function g(W, R) over the rectangular support $[\underline{r}, \bar{r}]^2$.[4] If the opportunity wage R^1 exceeds W^1, 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 $R^1 < W^1$, the employment relationship continues for another period at the new wage W^1. The wage rate is flexible, consequently separations are always efficient.

If growth is associated with equal improvement in worker productivity across the two firms (like accumulated general capital), then growth enters the model as a translation of the probability density function. That is, $g_\gamma(W, R) \equiv g(W-\gamma, R-\gamma)$ over $[\underline{r}+\gamma, \bar{r}+\gamma]^2$ where γ denotes growth between periods.[5]

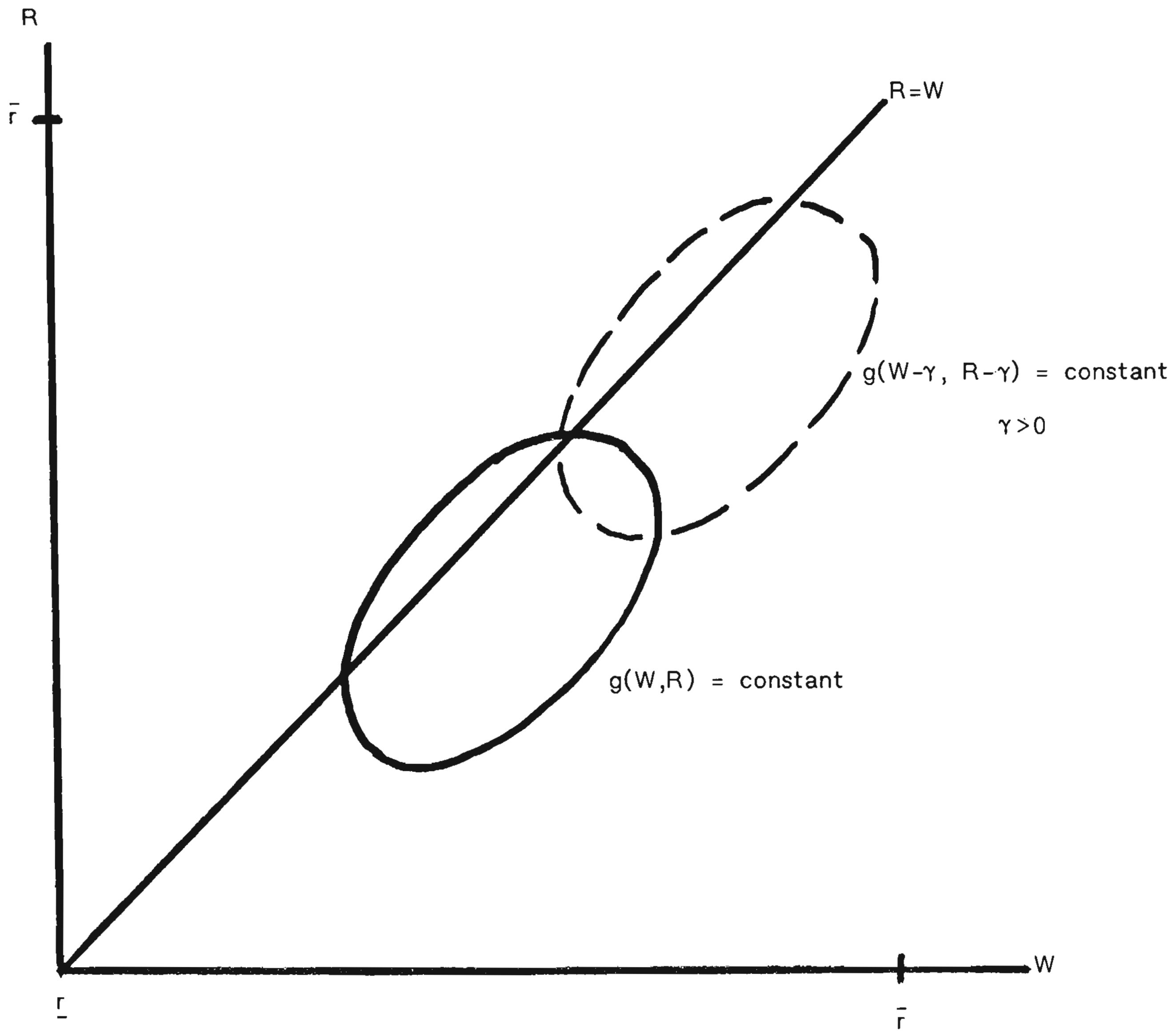
The model is summarized in Figure 1. To emphasize the stochastic feature of the model, a representative iso-probability contour is depicted. All draws in the half-space above the R=W ray result in separations. Growth in general skills corresponds to a northeast movement of the iso-probability contour in Figure 1. Accumulation of firm-specific skills, which corresponds to a rightward shift in the iso-probability contour, is not treated in this paper.

The separation rate s is just the probability of getting a random draw (W^1, R^1) in the half-space above the R=W ray, or the mass of the density in this region.

$$(2.0) \qquad s = \int_{\underline{r}+\gamma}^{\bar{r}+\gamma} \int_{W}^{\bar{r}+\gamma} g(W-\gamma, R-\gamma)\,dR\,dW = \int_{\underline{r}}^{\bar{r}} \int_{W}^{\bar{r}} g(W, R)\,dR\,dW,$$

which is independent of γ. (The equivalence of the two integral expressions is established by a change of variables.) The expected wage conditional on separation is[6]

Figure 1



$$(2.1) \qquad \bar{R}_s(\gamma) \equiv E[R|\gamma,\ S=1] = \int_{\underline{r}+\gamma}^{\bar{r}+\gamma} \int_{W}^{\bar{r}+\gamma} R\ g(W-\gamma,\ R-\gamma)/s\ dRdW$$

$$= \gamma + \int_{\underline{r}}^{\bar{r}} \int_{W}^{\bar{r}} R\ g(W,\ R)/s\ dRdW,$$

where the final expression is again established using a change of variables. Since $\Delta W = R - W^o$ for separations, the derivative of $\bar{R}_s$ with respect to γ measures the effect of general productivity growth on the expected wage growth of separations. The expected wage growth of separations moves one-for-one with growth in general productivity: $d\bar{R}_s/d\gamma = 1$.

The next step is to append to this model of efficient separations a theory of quits and layoffs. I consider several alternatives, and evaluate the success of each in matching up against the known regularities. For expositional convenience, I focus on the life-cycle regularities; thus experience E is the only interaction variable. However, the results also apply to any variable which effects the steepness of the life-cycle wage profile and to cyclical and secular variations in productivity.

Random Quit-Layoff Labels

The following simple scheme is easily rejected, but it is a useful introduction to the effect of a concave life-cycle productivity profile (i.e., γ declining with experience) on wage growth - turnover status interactions. A given separation is a quit with probability π and a layoff with probability $1-\pi$. The quit and layoff rates, πs and $(1-\pi)s$, are independent of the growth rate γ, implying also that their ratio $q/\ell = \pi/(1-\pi)$ is independent of γ. The expected wage for both quits and layoffs is $\bar{R}_s(\gamma)$; therefore, the expected wage gain to both quits and layoffs moves one-for-one with growth in general productivity.

This random labeling scheme accounts for neither the variation in the quit-layoff ratio over the life cycle nor differential wage growth by turnover type; nevertheless, it does explain why wage gains from separating fall with experience for both quits and layoffs. Wage gains by turnover type move one-for-one with general productivity growth which falls with experience. However, no interaction is implied when the direct effects captured by X_i are included in the empirical specification.

Quit-Layoff Labels and the Benchmark Wage

What determines whether a given separation is a quit or a layoff? It is not enough to assert that a quit is a worker leaving the firm and a layoff, the firm leaving or dismissing the worker. What determines "who leaves whom"? The common answer is the wage, but what wage? At some wage the firm would want to keep the worker, but the worker would choose to leave; however, at a sufficiently higher wage the firm would want to leave the worker, but the worker would prefer to stay. The standard solution in the literature on the quit-layoff distinction is to assume the wage is rigid, so quits and layoffs are determined relative to a fixed wage. In this approach, quits are voluntary separations and layoffs involuntary separations.

With efficient separations the voluntary-involuntary interpretation is vacuous since all separations are joint wealth maximizing.[7] One specification within the joint wealth maximizing approach applies quit-layoff labels based on a benchmark wage. For simplicity, I define a quit (layoff) to be a separation to employment at a wage exceeding (falling short of) the benchmark wage.[8] I consider two alternatives for the benchmark wage. The first is the pre-separation wage W^o augmented by general productivity growth γ: $W^o+\gamma$. The second does not augment W^o by γ, so it is just the pre-separation wage: W^o. This latter benchmark is employed in my other work on the quit-layoff distinction (McLaughlin 1985, 1986b).

In Figure 2, I illustrate the quit, layoff, and continued employment regions under the two benchmarks. The shaded region is a layoff under the growth-augmented benchmark, but a quit under the pre-separation wage benchmark. In terms of the empirical wage growth differentials of item (i), the growth-augmented benchmark implies quits have higher wage growth than layoffs. Positive wage growth to quits and negative to layoffs results from the pre-separation wage benchmark. While both benchmarks are consistent with this regularity, the pre-separation wage benchmark is more restrictive.

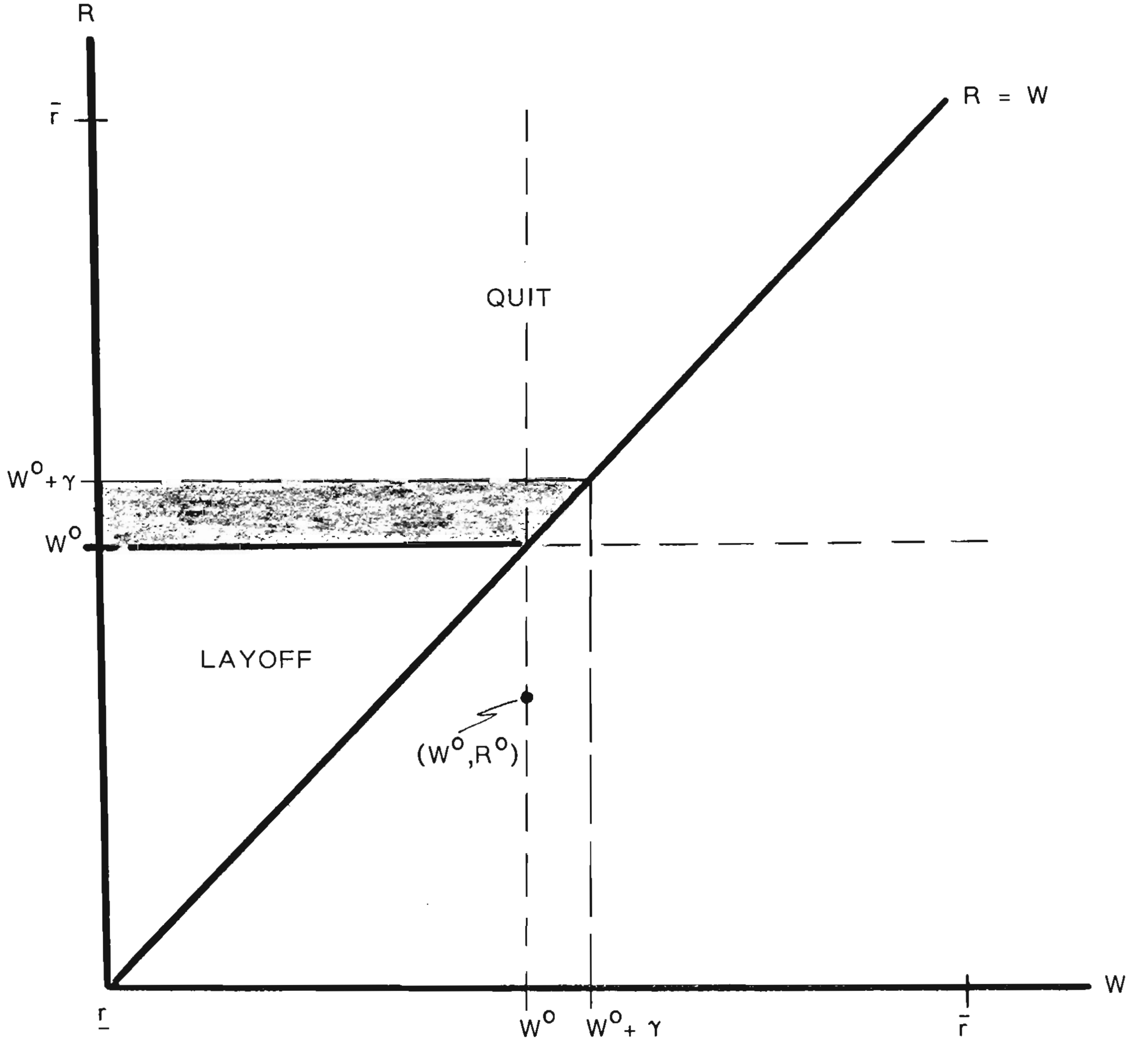
An implication that can be gleaned from Figure 2 is that unlike benchmark W^o, the growth-augmented benchmark is neutral with respect to growth γ. Under benchmark $W^o+\gamma$, variation in γ merely changes the units of measurement. The difference between the two benchmarks on this margin is sufficient to distinguish them empirically.

1. Growth-Augmented Benchmark: $W^o + \gamma$

The analytical representation of the quit and layoff rates under the growth-augmented benchmark is:

$$q(W^o) = \int_{W^o+\gamma}^{\bar{r}+\gamma} \int_{\underline{r}+\gamma}^{R} g(W-\gamma,\ R-\gamma)\,dWdR = \int_{W^o}^{\bar{r}} \int_{\underline{r}}^{R} g(W,\ R)\,dWdR \tag{3.0}$$

Figure 2



R
$\bar{r}$
R = W
QUIT
$W^o + \gamma$
W^o
LAYOFF
(W^o, R^o)
$\underline{r}$
W^o
$W^o + \gamma$
$\bar{r}$
W

$$(3.1) \qquad \ell(W^o) = \int_{\underline{r}+\gamma}^{W^o+\gamma} \int_{W}^{W^o+\gamma} g(W-\gamma,\ R-\gamma)\,dR\,dW = \int_{\underline{r}}^{W^o} \int_{W}^{W^o} g(W,\ R)\,dR\,dW.$$

Neither the quit nor the layoff rate is a function of general growth γ; hence this model of quits and layoffs does not yield the prediction of the quit-layoff ratio falling over the life cycle.

The expected wage by turnover type is:

$$(4.0) \qquad \bar{R}_q(W^o,\ \gamma) \equiv E[R|\gamma,\ Q=1] = \int_{W^o+\gamma}^{\bar{r}+\gamma} \int_{\underline{r}+\gamma}^{R} R\ g(W-\gamma,\ R-\gamma)/q(W^o)\ dW\,dR$$

$$= \gamma + \int_{W^o}^{\bar{r}} \int_{\underline{r}}^{R} R\ g(W,\ R)/q(W^o)\ dW\,dR$$

$$(4.1) \qquad \bar{R}_\ell(W^o,\ \gamma) \equiv E[R|\gamma,\ L=1] = \int_{\underline{r}+\gamma}^{W^o+\gamma} \int_{W}^{W^o+\gamma} R\ g(W-\gamma,\ R-\gamma)/\ell(W^o)\ dR\,dW$$

$$= \gamma + \int_{\underline{r}}^{W^o} \int_{W}^{W^o} R\ g(W,\ R)/\ell(W^o)\ dR\,dW.$$

Like the random quit-layoff labels, there is a one-for-one relationship between expected wage growth and general productivity growth γ for each turnover type: $\partial\bar{R}_q/\partial\gamma = \partial\bar{R}_\ell/\partial\gamma = 1$. Thus the current model accounts for the declining gains of separations in general, and of both quits and layoffs in particular, over the life cycle <u>only if</u> the X_i variables are omitted from the regression. No interaction is predicted. Therefore, the only advantage of this specification over the random scheme is in its implication of differential wage gains by turnover type.

2. Pre-Separation Wage Benchmark: W^o

Under the benchmark wage W^o, the analytical representation of the quit and layoff rates is

$$(5.0) \qquad q(W^o, \gamma) = \int_{W^o}^{\bar{r}+\gamma} \int_{\underline{r}+\gamma}^{R} g(W-\gamma, R-\gamma)\, dWdR$$

$$(5.1) \qquad \ell(W^o, \gamma) = \int_{\underline{r}+\gamma}^{W^o} \int_{W}^{W^o} g(W-\gamma, R-\gamma)\, dRdW.$$

Productivity growth γ translates the probability density function g northeast while preserving the turnover regions of the support. Thus growth moves part of the mass of the density out of the layoff region and into the quit region. Quits rise and layoffs fall by the same magnitude:

$$(6) \qquad \frac{\partial q}{\partial \gamma} = - \frac{\partial \ell}{\partial \gamma} = \int_{\underline{r}+\gamma}^{W^o} g(W-\gamma, W^o-\gamma)dW \equiv b(W^o, \gamma) \geq 0.$$

This is simply the probability of getting a draw at the boundary between the quit and layoff regions. Since quits are increasing and layofs decreasing in γ, and productivity growth is negatively related to experience, the quit-layoff ratio is predicted to fall over the life cycle.

The next result is that for neither quits nor layoffs does expected wage growth move one-for-one with general productivity growth. Both $\partial\bar{R}_q/\partial\gamma$ and $\partial\bar{R}_\ell/\partial\gamma$ are less than one. This is a selection result. Some low wage-growth separations (would-be layoffs) are pulled into the quit region reducing the positive effect of general productivity growth on the expected wage growth of quits. Of course, the layoff region is losing its highest wage growth separations to the quit region, thereby reducing the positive effect of γ for layoffs as well. Analytically, for quits

$$(7.0) \qquad \bar{R}_q(W^o, \gamma) = \int_{W^o}^{\bar{r}+\gamma} \int_{\underline{r}+\gamma}^{R} R\ g(W-\gamma,\ R-\gamma)/q(W^o,\ \gamma)\ dWdR,$$

and for layoffs

$$(7.1) \qquad \bar{R}_\ell(W^o, \gamma) = \int_{\underline{r}+\gamma}^{W^o} \int_{W}^{W^o} R\ g(W-\gamma,\ R-\gamma)/\ell(W^o,\ \gamma)\ dRdW.$$

These expressions imply the following two comparative static results:

$$(8.0) \qquad \frac{\partial \bar{R}_q}{\partial \gamma} = 1 - [\bar{R}_q(W^o,\ \gamma) - W^o]\ b(W^o,\ \gamma)/q(W^o,\ \gamma) \le 1$$

$$(8.1) \qquad \frac{\partial \bar{R}_\ell}{\partial \gamma} = 1 - [W^o - \bar{R}_\ell(W^o,\ \gamma)]\ b(W^o,\ \gamma)/\ell(W^o,\ \gamma) \le 1.$$

That each wage growth effect is positively related to its respective turnover rate implies interactions in regression (1.1) that decline with experience even with X_i included.[9] That is, $\frac{\partial^2 \bar{R}_q}{\partial \gamma \partial q}$ and $\frac{\partial^2 \bar{R}_\ell}{\partial \gamma \partial \ell}$ both positive implies that both $\frac{\partial^2 \Delta W}{\partial E \partial q}$ and $\frac{\partial^2 \Delta W}{\partial E \partial \ell}$ are negative in accordance with regularity (iii). This is established as follows for the case of quits. Let $\beta_E < 0$

measure the direct effect of experience on general productivity growth. Then

$$(9.0) \qquad \left.\frac{\partial\Delta W}{\partial E}\right|_Q = \left.\frac{\partial\Delta W}{\partial\gamma}\right|_Q \cdot \frac{\partial\gamma}{\partial E} = \frac{\partial\bar{R}_q}{\partial\gamma} \cdot \beta_E \leq 0;$$

therefore:

$$(9.1) \qquad \left.\frac{\partial^2\Delta W}{\partial E\partial q}\right|_Q = \frac{\partial^2\Delta\bar{R}_q}{\partial\gamma\partial q} \cdot \beta_E \leq 0$$

from equation (9.0).[10] The effect of experience on wage growth falls with the quit rate, or by the symmetry of cross-partials the gain to quitting falls with experience. The derivation for layoffs is entirely analogous.

Discussion

If the pre-separation wage W^o is the benchmark used in applying quit-layoff labels, the joint wealth maximizing approach to the quit-layoff distinction accounts for the life-cycle regularities. Two additional points extend the finding. First, the results apply to several other variables related to the growth rate of general productivity. Second, the rigid wage approach to the quit-layoff distinction fails to capture the empirical regularities.

In the previous section, I used the pre-separation wage benchmark in establishing two relationships. The first is that the ratio of quits to layoffs is increasing in the growth rate of general productivity γ. The second is that γ interacts with turnover status in a wage growth regression: wage growth of both quits and layoffs increase with γ less than one for one, and the magnitude of the "bias" depends on the respective quit and layoff rates. Labor market experience is one variable related to general productivity growth, consequently the model captures the quit-layoff regularities of the life cycle. Consider several other growth-related variables. In terms of aggregate cyclical variations, the ratio of quits to layoffs is predicted to vary procyclically; and business cycle variation is predicted to interact (positively) with turnover status in a wage growth regression. Since the steepness of the life-cycle profile is increasing in education (McLaughlin 1986a), the model predicts that the quit-layoff ratio is increasing in education and that, controlling for direct effects, the wage gains to both quits and layoffs are increasing in education. To the extent marriage, race, and sex affect the steepness of the life-cycle profile, the results apply to variables indicating such statuses as well.

Three variables which do not fit cleanly into the analysis are employment tenure, and indicators of industry of employment and union status. As modeled above, growth is purely general, improving outside opportunities by the same magnitude as productivity within the incumbent firm. For each of these three variables, this is not always true. The model clearly does not apply to employment tenure which improves opportunities with the incumbent employer relative to other firms. Thus the model does not predict an effect of employment tenure on the ratio of quits to layoffs, nor any interaction in the wage growth regression. If workers move across industries, then industry-specific growth rates cannot be treated as general growth. However, if most separations are to employers within the industry of the incumbent employer, perhaps little error is introduced by treating the "other firm" as a firm in the same industry, and hence treating industry-specific growth as general growth. With this qualification, the model predicts that the quit-layoff ratio covaries positively with the industry growth rate; indeed declining industries are predicted to have fewer quits and more layoffs than growing industries. Interactions between industry status and turnover type in the wage growth regression are also implied. Similarly for union status: If union workers separate to other "union employers," then a flatter life-cycle profile of union workers results in a lower ratio of quits to layoffs of union workers, and (negative) interactions between union status and turnover-type in the wage growth regression.[11]

A second issue is whether the more standard approach based on wage rigidity captures the growth-related regularities. With the exception of my work, the current models of the quit-layoff distinction rely on rigid wages. (See, e.g., Hashimoto and Yu (1980), Hall and Lazear (1984), and my survey of the rigid wage approach (McLaughlin 1986a).) If the wage in a rigid wage model is fixed, then the results are quite similar to those of the joint wealth maximizing model under the pre-separation wage benchmark.[12] However, the similarity is illusory since the wage would not be fixed at W^o even in a rigid wage model.

Consider briefly a property of an optimal contract between a risk neutral firm and a risk neutral worker under asymmetric information. The wage in the optimal contract is inflexible with respect to productivity shocks which are not observable to both the firm and the worker. However, the wage is flexible with respect to those components of productivity variation which are forecastable, proxied by observables, or verifiable once revealed by either party to the contract. General productivity growth is a component of productivity variation which induces wage flexibility. Consider two examples: Even in the rigid wage model, the wage is flexible with respect to the life-cycle evolution of worker skills since such skill accumulation is forecastable. Over the life cycle the conventional approach predicts a rigid wage profile, not a rigid wage. The wage is also flexible over the business cylce since the stage of the business cycle, if not directly verifiable, is easily proxied by observable variables.

Since the growth-related movements under study here would be written into the optimal contract, the wage is predicted to be flexible on this margin. Consequently, in terms of accounting for the empirical regularities, the "rigid wage" model (at best) fares similarly to the efficient separations model with the growth-augmented benchmark: neither accounts for the regularities in item (iv) or wage-growth - turnover-type interactions.

Pulling these results together, I conclude that the benchmark I have employed in previous analysis of the quit-layoff distinction can account for the several empirical regularities listed in the introduction. The life-cycle pattern of quits and layoffs, and the interaction effects of experience and turnover type in wage growth regressions are captured by this specification. In addition, results extend to any--indeed all--variables related to general productivity growth. These include education and measures of the business cycle. Finally, the empirical regularities do not support the rigid wage model. The rigid wage model predicts comovements only for those productivity variations which cannot be contracted over in advance.

IV. DIRECT EVIDENCE

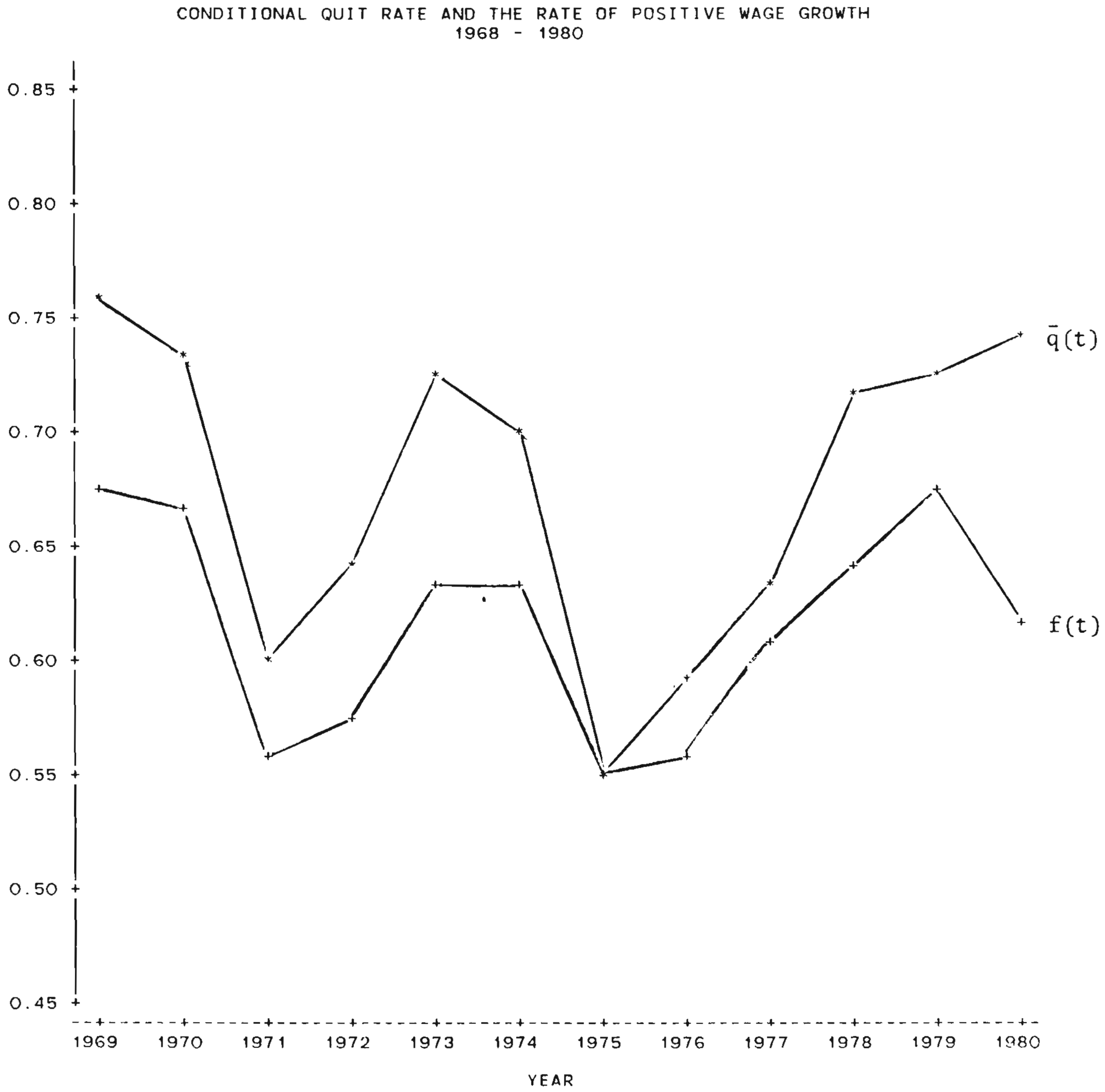
Is there direct evidence to support the pre-separation wage benchmark in particular and the the joint wealth maximizing approach in general? The theory predicts that the fraction of separations labeled quits is the fraction of separations accepting wages exceeding the benchmark, here W^o. Thus an exact predictor of $\bar{q}$ the quit rate conditional on separation is f the fraction of separations satisfying $R > W^o$. I term f the rate of positive wage growth.

I employ data from the Panel Study of Income Dynamics over the years 1968-1980 to investigate the efficacy of this prediction. Since I am concerned with employment transitions, the data are pooled into twelve annual transitions; that is, 1968-1969, 1969-1970, ..., 1979-1980. (I refer to the first year of each pair as the initial period.) For each observation, the data include information on individual and employment characteristics over both years of each transition. In particular, each respondent who changes employers in the intervening year is asked the reason for leaving the job held at the initial period's interview, and whether the new job pays more. Using these two variables, I compute indicator variables $\bar{Q}$ and F which indicate whether the separation is a quit and whether the new job pays more. The sample is limited to male household heads aged 18 to 58 in the initial period, who are initially employed, and who separate and describe the separation as either a quit or a layoff. The sample consists of 3,310 underlying observations which are grouped by year (about 275 observations per year).

To evaluate the prediction, I focus on two margins: whether the prediction f is accurate in terms of the level of conditional quits $\bar{q}$, and second the comovement between f and $\bar{q}$ over time. Figure 3 illustrates both the levels of and comovement between the conditional quit rate and the rate of positive wage growth. While f clearly lies below $\bar{q}$, the level is approximately right: the sample mean of $\bar{Q}$ is .68 and of F is .62. There is strong contemporaneous covariation over the time period; it appears from Figure 3 that the two series move close to one-for-one.

Although the prediction of the model is borne out by the cursory evidence, it does not survive formal testing. Consider the test of equality of the two means. The means (standard deviations) of $\bar{Q}$ and F are .680 (.4665) and .617 (.4861). Assuming the normal approximation to the binomial

FIGURE 3


CONDITIONAL QUIT RATE AND THE RATE OF POSITIVE WAGE GROWTH
1968 - 1980
0.85
0.80
0.75
0.70
0.65
0.60
0.55
0.50
0.45
1969
1970
1971
1972
1973
1974
1975
1976
1977
1978
1979
1980
YEAR
$\bar{q}(t)$
$f(t)$

distribution, the t-statistic associated with this test is 5.38, which rejects the null hypothesis of equality of the two means.

To analyze the comovement between the conditional quit rate and the fraction of separations with positive wage growth in the employment transition, I develop the following regression specification. Let $I \equiv a + bF + \epsilon$ denote an index function with $F = 1$ if $R^1 > W^0$, and $F = 0$ otherwise; ϵ is a random variable capturing errors in reporting or other determinants of the quit-layoff distinction such that $E\epsilon=0$ and $E[\epsilon F]=0$. Then the conditional quit rate is expressed: $\bar{q}(F) \equiv \Pr[\bar{Q}=1 | F] = \Pr[I>0 | F] = \Pr[\epsilon > -(a + bF)]$. If $\epsilon \sim N(0, \sigma^2)$, then $\bar{q} = \Phi(a^* + b^*F)$ where Φ is the standard normal cumulative distribution function and $a^*=a/\sigma$ and $b^*=b/\sigma$ are the probit coefficients. The theory implies the test of the null hypothesis of $a=0$ and $b=1$. However, in the probit setting neither a nor b is identified; in particular one cannot test the hypothesis $b=1$. An alternative is to assume ϵ is uniform over the interval $[0, 1]$. Hence, $\bar{q}(F) = a + bF$. This is a linear probability model.

The results of minimum χ^2 estimation of the linear probability model on grouped data are reported in Table 1.[13] (Maximum Likelihood estimates of the probit model are presented for reference.) Separate t-tests of zero intercept and unit slope fail to reject the null hypotheses implied by the theory. However an F-test of the joint hypothesis clearly rejects the model's prediction. Figure 4, a scatter diagram of the data, illustrates the nature of the rejection. To support the joint wealth maximizing approach, the data must lie along the 45° line. Clearly, the data are clustered above the 45° line. While a zero intercept or a unit slope can be imposed separately with little loss, simultaneous imposition of the two is too restrictive.

I conclude that the model does reasonably well in predicting the fraction of separations labeled quits; in particular, time series variation in the

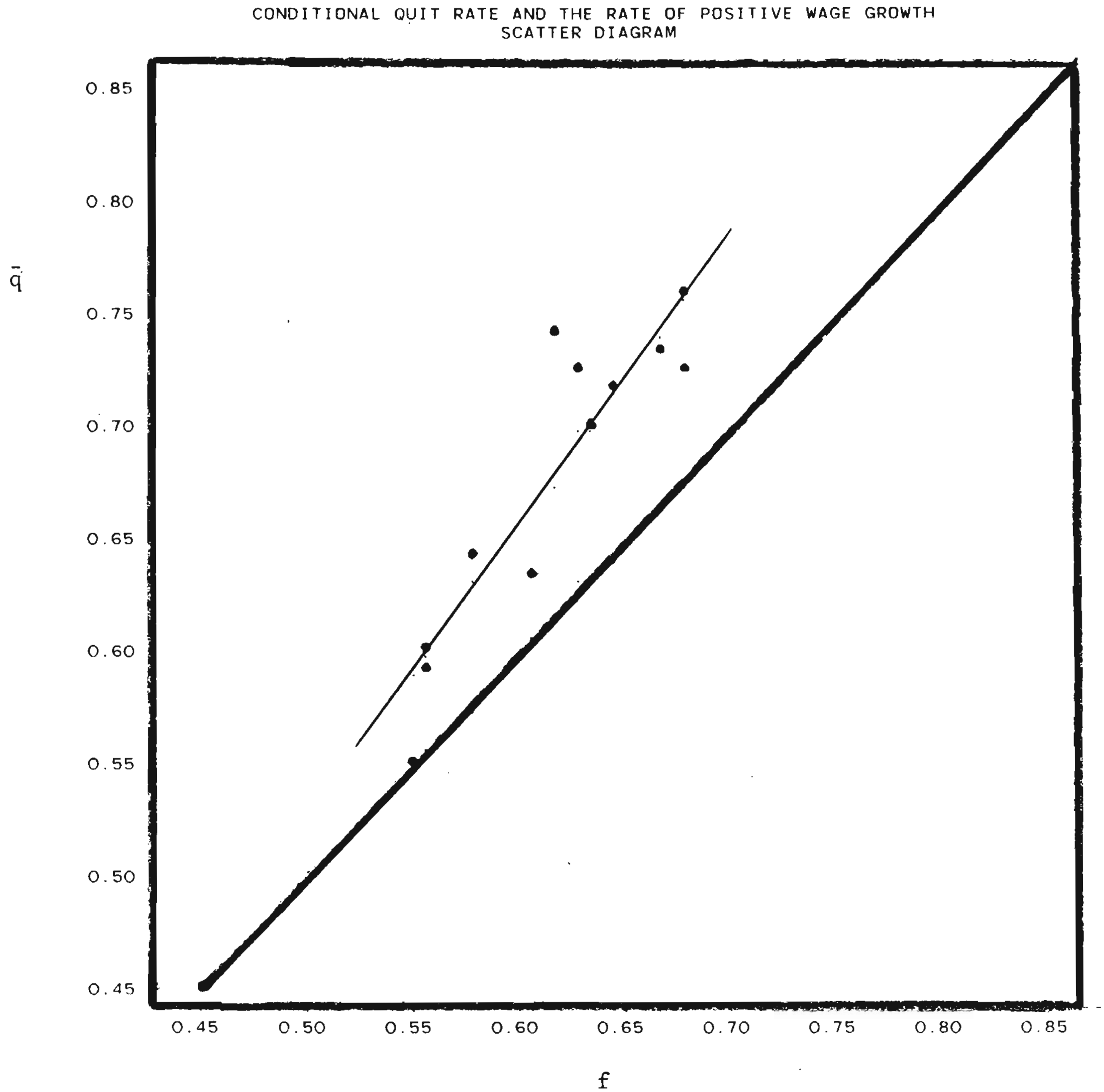
TABLE 1

CONDITIONAL QUIT REGRESSION

Model (Estimator)	constant	F	R^2	DW	F-Stat
Linear Probability (Minimum Chi-Square)	-0.105 (0.134)	1.271 (0.216)	.78	1.77	21.3
Probit (Maximum Likelihood)	-1.737 (0.320)	3.581 (0.519)	.87	-	-

Note: asymptotic standard errors in parentheses. 12 observations per regression from 3,310 underlying observations. The R^2 for the probit regression is from the (unreported) first-stage probit minimum-chi-square estimates. The F-statistic corresponds to the joint test of a=0 and b=0 (see text), and thus has 2 restrictions and 10 degrees of freedom.

FIGURE 4


CONDITIONAL QUIT RATE AND THE RATE OF POSITIVE WAGE GROWTH
SCATTER DIAGRAM
0.85
0.80
0.75
0.70
0.65
0.60
0.55
0.50
0.45
$\bar{q}$
0.45
0.50
0.55
0.60
0.65
0.70
0.75
0.80
0.85
f

conditional quit rate is closely approximated by time series variation in the fraction of workers separating to higher paying jobs. However, since a formal test rejects the restrictions implied by the model, my specification is not a complete characterization of labor market turnover.

VI. CONCLUSION

In this paper, I confront the problem of interpreting and estimating the effects of turnover on wage growth. Although I do not dismiss Mincer's (1986) interpretation based on the arrival rate of wage offers, I develop a refutable alternative which accounts for the growth-related empirical regularities by a reverse causation: wage growth determines whether a given employment transition is a quit or layoff.

The joint wealth maximizing approach to labor turnover is sufficiently flexible to admit a variety of specifications. Nevertheless, the particular specification which I adopt--based on the pre-separation wage benchmark-- is quite restrictive. The model implies all variables related to general productivity growth interact with turnover status in the wage growth regression; and the ratio of quits to layoffs varies with all such variables as well.

The empirical results in section V are mixed. While the structural comovements implied by the model are clearly present in the data, the restrictions are formally rejected.

Although the <u>structural</u> estimates indicate that my specification is not a complete characterization of the quit-layoff distinction, the joint wealth maximizing approach clearly goes a long way in accounting for many features of labor turnover. Indeed, the model improves on the current literature on the quit-layoff distinction because the rigid wage model is deficient in capturing the growth-related empirical regularities.

APPENDIX

ECONOMETRIC SPECIFICATION

Two approaches have been taken in the empirical literature on labor turnover. The first, the estimation of wage growth regressions by turnover status, is the central theme of this paper. Examples of this work can be found in Bartel and Borjas (1981), Antel (1985), and Mincer (1986). The second focuses on the determinants of turnover status, in particular the effect of potential wage growth on separation, quit, and layoff rates. Borjas and Rosen (1980) and McLaughlin (1985b) are examples of this approach. My model links these two approaches by showing how endogenous quit-layoff labels affect the parameter estimates in wage growth regressions. Consider the following simultaneous equations specification:

$$\text{(A.0)} \qquad \Delta W_i = \alpha + X_i\boldsymbol{\beta} + \delta_Q Q_i + \delta_L L_i + (Q_i G_i)\delta_Q^G + (L_i G_i)\delta_L^G + \epsilon_i;$$

$$\text{(A.1)} \qquad \Delta W_i^* \equiv R_i - W_i^o = \alpha + X_i\boldsymbol{\beta} + \epsilon_i;$$

$$\text{(A.2)} \qquad S_i^* \equiv R_i - W_i = Z_i\boldsymbol{\gamma} + u_i;$$

$$\text{(A.3)} \qquad Q_i = 1 \quad \text{if} \quad S_i^* > 0 \;\wedge\; \Delta W_i^* > 0, \qquad = 0 \text{ otherwise};$$

$$\text{(A.4)} \qquad L_i = 1 \quad \text{if} \quad S_i^* > 0 \;\wedge\; \Delta W_i^* \le 0, \qquad = 0 \text{ otherwise}.$$

This is an example of the dummy endogenous variables model termed by Heckman "the hybrid model with structural shift," extended to multivariate criteria for the dummy variables (Heckman 1978).[14] Although the restrictions of my hypothesis (i.e., $\delta_Q = \delta_L = 0$, and $\delta_Q^G = \delta_L^G = 0$ for all variables in G) are not imposed on the wage growth equation (A.0), my turnover model is imposed in equations (A.1) - (A.4) for the determination of the dummy variables.

The dummy variables in equations (A.0) - (A.4) play two roles. Q_i and L_i in (A.3) and (A.4) are indicators of the latent variables S_i^* and ΔW_i^* crossing thresholds. However, in equation (A.0), Q_i and L_i are structural shift variables; that is, variables which shift the behavioral relationship. The estimation problem is to identify these two roles. If solved, one can distinguish between my approach, which limits the role of Q_i and L_i to indicators of the latent variables, and structural shift interpretations such as in Mincer (1986).

The properties of OLS estimates of equation (A.0) have been discussed informally above in the context of the theoretical model. Heckman (1978) presents the formal treatment in a more general context, and proposes several estimators with desirable properties. I limit my discussion to the instrumental variables (IV) estimator in the presence of nonlinearities in the endogenous variables.

In particular, I am interested in identifying the parameters of the linearized version of equations (A.0) - (A.4):

$$\text{(A.0)} \qquad \Delta W_i = \alpha + X_i\beta + \delta_Q Q_i + \delta_L L_i + (Q_i G_i)\delta_Q^G + (L_i G_i)\delta_L^G + \epsilon_i;$$

$$\text{(A.3')} \qquad Q_i = X_i\eta_Q^x + Z_i\eta_Q^z + \epsilon_{Qi};$$

$$\text{(A.4')} \qquad L_i = X_i\eta_L^x + Z_i\eta_L^z + \epsilon_{Li}.$$

Thus (A.1) - (A.4) are replaced by linear probability models for Q_i and L_i. (Structural estimation of (A.1) - (A.4) is analyzed in McLaughlin (1986b).) Since the X_i and Z_i variables are taken as exogenous, the issue of identification focuses on the wage growth regression (A.0).

Although linear in the parameters, the wage growth regression is nonlinear in the endogenous variables due to the turnover-type interactions. (Fisher (1966) analyzes identification in this class of models.) The wage growth equation is identified if and only if Z_i contains at least two variables distinct from the set of growth variables X_i.

The theory does predict that X_i and Z_i are not identical. In particular, employment tenure is a variable in the set X_i which is excluded from Z_i. However, this does not aid in the identification of the parameters of (A.0). The theory also predicts <u>one</u> variable in Z_i is excluded from X_i: the pre-separation wage W_i^o. In the absence of further exclusions the wage growth regression is fundamentally under-identified. Nevertheless, one variable from outside the theoretical model can be employed to identify the system: subsidies to unemployment insurance benefits. In McLaughlin (1986b), I briefly discuss the incentives to "relabel" quits as layoffs in the presence of (subsidized) unemployment insurance benefits. The degree of subsidization of unemployment insurance benefits decreases quits relative to layoffs, reduces total separations, but has no effect on wage growth. Thus the parameters of the wage growth equation can be identified.

NOTES

* I thank John Boyd, Barbara Mace, Tom Mroz, and workshop participants at the University of Chicago and the University of Rochester for helpful comments.

1. It is important to realize that regression (1.1) is not a differenced wage regression--the regressors are not time differenced--but is indeed a wage growth regression.

2. I have three additional remarks on this point. First, the reservation wage of the on-the-job searcher is unchanged. Second, there will be fewer quits due to fewer on-the-job searchers as well. Some workers will switch from working and searching to just working. Third, Mincer neglects the result from OJS theory that quits decline with age in the absence of reduced arrival rate of wage offers (Burdett 1978). Simply, older workers have had more "draws," implying a superior expected maximum draw (a state variable), and thus a lower probability of finding a higher wage offer.

3. In this paper, I limit the analysis to the case of two firms. See McLaughlin (1986b) for the generalization to n firms. In addition, I assume the worker is paid his productivity value and thus captures any rents associated with the match. In McLaughlin (1987), I analyze the effect of rent sharing on turnover.

4. In general, g(W, R) depends on the identity of the incumbent employer. For notational ease I suppress this element.

5. From this point forward in the text, I use the (W, R) notation without any reference to the particular period or γ. Therefore, explicitly, let g be the bivariate probability density function associated with the random variables (W^0, R^0); and let $W^1=W^0+\gamma$ and $R^1=R^0+\gamma$. Then the bivariate probability density function of (W^1, R^1) conditional on γ is

$$g_\gamma(W^1, R^1) \equiv g(W^1-\gamma, R^1-\gamma),$$

as stated in the text without the superscripts. At several points in the text below results are established using this change of variables.

6. The presence of s in the denominator of the integrand is required to make the density integrate to one. That is, the operation requires integration over the conditional density.

7. The appropriate interpretation in this context is that quits are worker-initiated and layoffs firm-initiated separations; initiations of would-be inefficient separations are counteracted by wage flexibility or side payments.

8. In McLaughlin (1986b), I offer a more detailed analysis of the process governing the quit-layoff labels, and link "initiations" to the benchmark wage. The definition in the text is an implication in the more detailed analysis.

9. However, an interaction is not predicted if the coefficients are restricted to analyze the effects of separations. See the discussion following equation (2.1).

10. I have ignored the effect of the quit rate on the term $(\bar{R}_q - W^o)b$. The effect is distribution specific and in general ambiguous.

11. A more satisfactory approach to the effect of union status on quits and layoffs is developed in McLaughlin (1987). There I relax the assumption that workers capture all the match rents, and investigate the effect of rent sharing on quits and layoffs. I conclude that if union workers capture a higher share of the match rents than their nonunion counterparts, then union workers should exhibit a higher quit-layoff ratio, and the usual interactions should also apply.

12. The results of the two approaches are not identical. Nevertheless, in terms of the comparative statics of productivity growth, the differences are higher order.

13. The problems with the linear probability model--biased and inconsistent estimates, and predicted "probabilities" outside the unit interval-- are less severe when estimated on group data. See, e.g., Maddala (1983, 28-29).

14. If the empirical work is limited to mover-stayer comparisons, the model reduces to a form identical to that analyzed in Heckman (1978).

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