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In the labor market, the "labeling" of employment transitions as quits or layoffs is pervasive. Why? Is a quit different than a layoff? If quits and layoffs do appear to be different, is there any economic content to the terminology or is there no meaningful economic distinction?

Previous research has approached the quit-layoff distinction with a voluntary-involuntary interpretation. Under this interpretation, quits are voluntary to the worker (but might be involuntary to the firm) and layoffs are involuntary to the worker but voluntary to the firm. Economic models consistent with the voluntary-involuntary interpretation have relied on wage rigidity and the associated economic inefficiency.[1] Becker's (1962) early work on firm-specific human capital generated a distinction between quits and layoffs as a result of the sharing rule governing the costs of and returns to firm-specific investments. Since the sharing rule is a rigid wage profile, the wage revisions or side payments which would undermine the voluntary-involuntary interpretation are precluded. More recently, forms of wage rigidity have been introduced explicitly and supported by informational asymmetries and costly renegotiation (Hashimoto and Yu 1980; Hall and Lazear 1984; and Antel 1985). Models of on-the-job search (Parsons 1973 and 1977; Burdett 1978) yield implications for quit behavior for a given wage on the current job. Consequently, a common feature of these models is that, relative to the worker, quits are voluntary separations and layoffs involuntary separations.

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[1]See McLaughlin (1987b, Chapter 1) for a survey of the current literature on the quit-layoff distinction.

As an alternative, a model of "Who Leaves Whom" (McLaughlin 1987b, Chapter 2) can be applied to this problem to determine whether quits and layoffs are <u>economically</u> distinct employment transitions. If the Who-Leaves-Whom model applies, then the quit-layoff distinction is economically inconsequential--that is, labor is allocated across firms efficiently and turnover is joint wealth maximizing. But the quit-layoff labels are applied systematically, generating the observed behavioral differences.[2] In short, the quit or layoff label ends up <u>summarizing</u> the labor market environment rather than determining it.

My specification of the joint wealth maximizing approach includes the following features:

(i) A simple matching framework.

(ii) Employment relationships form in spot markets with flexible wages.

(iii) All separations are efficient, so the voluntary-involuntary interpretation does not apply.

(iv) Turnover labels (quit or layoff) are applied based on who initiates the separation by demanding a wage revision.

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[2] That quits and layoffs are behaviorally distinct is well known. The empirical regularities include, but are not limited to:

(i) Quits (layoffs) exhibit higher (lower) wage growth in the employment transition than those who do not separate.

(ii) Quits (layoffs) have a lower (higher) probability of experiencing an intervening spell of unemployment.

(iii) Quits are procyclical, layoffs countercyclical.

(iv) The ratio of quits to layoffs declines with age or experience.

See McLaughlin (1987b, Chapter 4) for documentation of these and several other empirical regularities of the quit-layoff distinction.

The quit-layoff distinction arises out of a censoring of wage revisions. Initial asymmetries in the information structure create private incentives for revising the wage. If a separation results, the side (firm or worker) which initiates the wage revision determines the turnover label.

Beyond the empirical regularities listed in note 2 above, the model yields such structural implications as the ratio of quits to layoffs is decreasing in the pre-separation wage and increasing in the value of outside opportunities. Effects on separations and the turnover labels of education, experience, tenure, union status, race, and cyclical fluctuations are produced in the reduced form. A side hypothesis considers the effect of subsidized unemployment insurance in increasing the incidence of layoffs relative to quits.

The paper is organized into five parts. In the first section, I set up the framework for subsequent analysis. The behavior of firms and workers in initiating wage revisions is used to define the turnover labels. Details of the matching model and labeling process are presented and comparative static results derived. The reduced-form implications are also discussed. In preparation for estimation and testing, the theoretical model is recast in section 2 as an empirical specification. I show that the model is estimable as a bivariate probit with selection (one empty cell) or as two univariate probits. In the latter implementation, one probit corresponds to the separation decision and the second to the turnover labels conditional on separation. In section 3, I discuss the data--a sample drawn from the Panel Study of Income Dynamics--employed to test the model's refutable implications. The empirical results, which are mixed, are discussed in section 4. The final section contains a summary and the conclusions.

1. Matching Model with Quits and Layoffs

The approach I take in analyzing the quit-layoff aspect of labor turnover follows the suggestion of Becker, Landes, and Michael (1977) in their analysis of turnover in the marriage market: labor turnover is always efficient or joint wealth maximizing, and as a result the voluntary-involuntary interpretation of turnover is inappropriate. The firm and worker dissolve their employment relationship if and only if their total value when separated exceeds the combined value of the match. Potentially inefficient separations result in wage adjustments or side payments to preserve the optimal assignment.

Assuming the wage rate or the division of the match value is perfectly flexible, it is not meaningful to say that the worker walked out and the firm was abandoned, or vice versa. The notion of voluntary versus involuntary turnover is lost. 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 the worker to leave, but the worker would prefer to stay.

If this is the appropriate approach to labor turnover, then what generates the behavioral differences between quits and layoffs? My thesis, following Becker, Landes, and Michael's (1977, 1145, n. 4) suggestion, is that labels of quits and layoffs are applied relative to the pre-separation division of the match value. The pre-separation wage is the benchmark used to divide separations into quits and layoffs. It is established below in the context of the model that the pre-separation wage is empirically distinguishable from other potential benchmarks and is supported in its predictions of the several empirical regularities.

The joint wealth maximizing hypothesis is an approach, not a formal specification. In this section, I formalize the approach in developing a set of refutable implications regarding the determinants of labor turnover. The specification is a simple matching model augmented with the turnover labels--quits and layoffs--but otherwise stripped down to its barest essentials.

Summary of the Model

One of the essential characteristics 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 wage W^0, and this worker has an opportunity wage with other firms of R^0, and a fixed nonmarket value R_0. (Superscripts denote the period, subscripts the sector or firm.) Between periods idiosyncratic shocks, which become common knowledge in the wage revision process, arrive changing the match values. If the new reservation value R^1 exceeds W^1, the new value of employment with the incumbent firm, the firm and worker dissolve the employment relationship and form new matches with their best alternatives. Alternatively, if the worker's value to the incumbent firm exceeds his new reservation value, the employment relationship continues for another period at the new wage W^1. If both W^1 and R^1 are less than R_0, the worker separates to the nonmarket sector. The wage rate is flexible and separations are always efficient.

In this model, what distinguishes quits from layoffs? Based on the thesis introduced above, quits fall into the following scenario. The worker receives an unusually good draw for productivity (wages) outside the incumbent firm: $R^1 > W^0$. The firm does not directly observe the draw, but it can be verified costlessly once the worker reveals the value to the firm. The worker asks his employer to meet or better it. If $W^1 < R^1$, the firm declines and the worker quits. Quits are worker initiated. Layoffs fall into a symmetric scenario. The firm sees that the worker's productivity drops, but this is unknown to the worker. The firm asks the worker to accept a wage cut of $W^0 - W^1 > 0$, and the worker verifies the productivity loss. If W^1 is less than R^1 or R_0, the worker does not accept the wage cut and he is laid off. Layoffs are firm initiated. A principal result of the two sides' behavior in initiating wage changes is that quits leave for better opportunities, layoffs for inferior opportunities.[3] This result establishes the basic distinction between quits and layoffs, and is consistent with the evidence.[4]

The model is summarized in Figure 1. To emphasize the stochastic feature of the model, a representative iso-probability contour is depicted. All draws

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[3] Becker, Landes, and Michael (1977, 1145, n. 4) also suggest that "a more promising approach relies on the cause of a job or marital separation. A quit could be said to result from an improvement in opportunities elsewhere and a layoff from a (usually unexpected) worsening in opportunities in this job or marriage." Borjas and Rosen (1980, 161-63) also discuss such labels, but do not incorporate the labeling process into their estimation of the model. They estimate a structural model of separation behavior, maintaining the hypothesis of no meaningful distinction between quits and layoffs.

[4] Bartel and Borjas (1981, Tables 2.1 and 2.2), Antel (1985, Table 3), Mincer (1986, Tables 1 and 1A), and McLaughlin (1987b, Tables 10 and 11) present evidence of differential wage growth in employment transitions by turnover type. Wage growth between jobs is higher for quits than for layoffs. The hypotheses of positive wage growth for quits and negative for layoffs is tested in McLaughlin (1987b, Chapter 4). The results indicate that quits exhibit positive wage growth, but one cannot reject the null hypothesis of zero wage growth of layoffs.

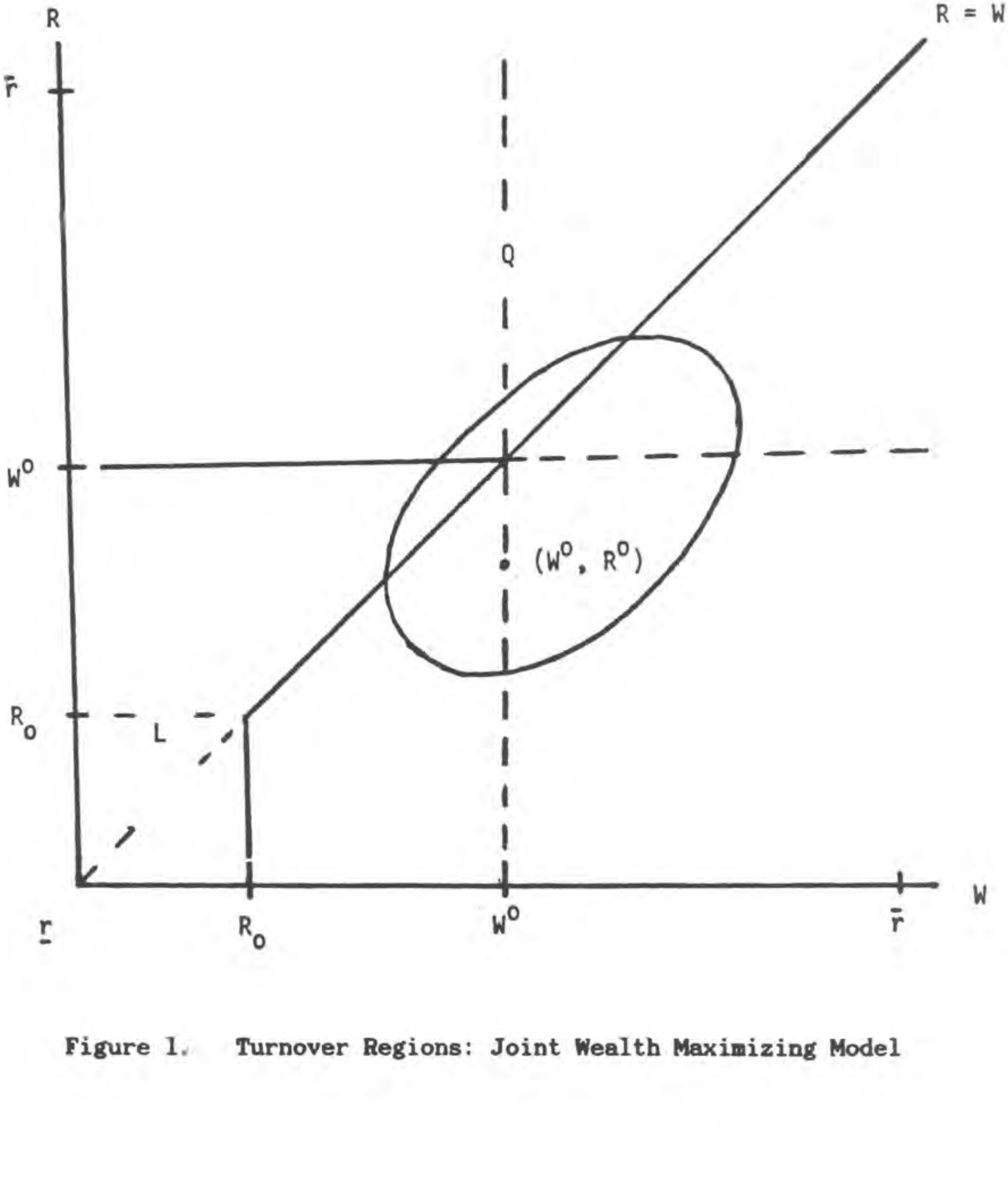

Figure 1. Turnover Regions: Joint Wealth Maximizing Model

in the half-space below the $R = W$ ray satisfying $W > R_0$ result in continued employment. Based on the initial draw (W^0, R^0) and the value of nonmarket production R_0, 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 various regions.

Structure of the Model

In this section, I extend the model to include n firms and present the details of the matching model and labeling process. The latter includes a discussion of the informational assumptions, and a formal definition of "initiations" which determines the turnover labels. I also reduce the dimensionality of the problem for tractability, and present the principal comparative static results. One of the costs of this subsection's generality is a complicated notation; however, the (W, R) notation reappears through the results of this subsection.

The representative, though heterogeneous, worker can match with any one of n firms. The value of the worker's productivity with firm i is denoted r_i.[5] Let $\mathbf{r}$ denote the n-vector of productivity values $(r_1, \ldots, r_n)$. The n-vector of wage offers $\mathbf{R}$ is in general related to $\mathbf{r}$. Assume that the worker receives any rents associated with a match, thus $\mathbf{R} = \mathbf{r}$. (This assumption is relaxed below.)

Separations result from unsuccessful attempts to change the wage payment, and whether a quit or layoff label is applied to the separation is determined

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[5]See McLaughlin (1987a) for an equilibrium analysis of the content of "productivity" in the matching context.

by the identity of the side initiating the wage change. If the firm does not accept the worker's demand for a wage increase, the worker quits. Likewise, if the worker refuses a firm-initiated wage cut, the worker is laid off.

The process of wage revision depends critically on the information structure. My basic assumption is that an initial informational asymmetry is resolved in the wage revision process. Consider the case of a worker employed with firm j in the previous period, so j is the incumbent employer. Stochastic shocks arrive changing the worker's productivities r. The worker is assumed to know all his outside offers, but is unaware of his productivity with firm j. Firm j knows only r_j. A second assumption is that all private information is costlessly verifiable once revealed (Grossman 1981; Milgrom and Roberts 1986). Therefore, there are no gains to the worker in generating fraudulent outside wage offers, nor to the firm in misrepresenting the worker's productivity.

Under these circumstances the worker has an incentive to initiate a wage increase if his best outside offer dominates his pre-revision wage; that is, if there exists some $R_i > W^o$ ($\equiv R_j^o$) for $i \neq j$. The firm accepts the wage increase if, for $i \neq j$, $R_i < R_j$ ($\equiv r_j$). Otherwise, the firm rejects the wage increase and the worker quits. <u>Quits are separations resulting from censored wage increases</u>. The layoff label is produced similarly. The firm has an incentive to initiate a wage cut if $r_j < W^o$. That is, if the worker's current value of productivity is less than his pre-revision wage, the firm initiates a wage cut. If $R_i > R_j$ ($\equiv r_j$) for some $i \neq j$, then the worker does not accept the wage cut and is laid off. <u>Layoffs are separations resulting from censored wage cuts</u>.

The initiation rules are not mutually exclusive. There is a set of possible draws in which both the worker and the firm initiate a wage change, but in opposite directions. Here I call the resulting separation a quit.[6]

It is useful to introduce a nonmarket or household sector. The worker values nonparticipation at R_0, which is nonrandom. The worker leaves the market sector if $R_0 > R_i$, for all $i = 1, \ldots, n$. Since the worker starts out employed in firm j, it must be the case that $R_0 < W^o$. Therefore, separations to the nonmarket sector are never labeled quits.[7]

Any particular turnover rate is given by the probability of getting a draw in a specified region of the productivity space. The quit and layoff regions, which depend on the identity of the incumbent employer, are given by the sets:

$$(1.1) \qquad Q(W^o;j) = \{R \in [\underline{r}, \bar{r}]^n: \ R_i \geq W^o \ \wedge \ R_i \geq R_j, \ \text{ for all } i \neq j\}$$

$$(1.2) \qquad L(W^o,R_0;j) = \{R \in [\underline{r}, \bar{r}]^n: \ W^o \geq R_i \geq R_j \ \text{ or } \ R_0 \geq R_j \geq R_i, \ \text{ for all } i \neq j\}.$$

The R (or r) are random variables assumed to be distributed independently, but

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[6]Three alternative treatments of these other separations are: (1) random allocation to Q and L, (2) allocation depending on the closeness of the draw to the Q and L regions, and (3) introduction of a third category, "mutual separations." Adoption of any one of these alternatives would not affect the principal qualitative results derived below.

[7]Letting R_0 be random adds little to the analysis. The only advantage is that some quits exit the market sector, but in smaller numbers than layoffs.

not necessarily identically,[8] so the joint density of R is

$$(2) \qquad g(R) = \prod_{i=1}^{n} g_i(R_i)$$

over the n-dimensional rectangular support $[\underline{r}, \bar{r}]^n$. The associated quit and layoff rates, or transition probabilities, are

$$(3.1) \qquad q(W^o;j) = \int_{Q(W^o;j)} g(R)dR$$

$$(3.2) \qquad \ell(W^o,R_o;j) = \int_{L(W^o,R_o;j)} g(R)dR$$

which depend on the identity of the incumbent employer. Comparative static calculations require differentiation through the regions of integration. By reducing the dimensionality of the problem, explicit solutions can be derived. This requires the use of order statistics.

Let R denote the best opportunity in the n-1 other firms. The density of the (n-1)*th* order statistic R of the n-1 independent but not necessarily

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[8]In writing equation (2), I am implicitly conditioning on market-wide shocks. Hence the remaining source of randomness is firm specific. See McLaughlin (1987b, Chapters 5 and 6) for my treatment of secular and cyclical market-wide shocks. Also note that I require independence only for tractability in writing the order statistic. Independence is not an essential characteristic of the model.

identically distributed random variables is[9]

$$(4) \qquad \bar{g}_j(R) = \sum_{k \neq j} \{g_k(R) \prod_{i \neq k} G_i(R)\}.$$

The joint density of the productivity in the incumbent firm (defining $W \equiv R_j$) and the best productivity in the n-1 other firms is

$$(5) \qquad g_j(W, R) = g_j(W)\, \bar{g}_j(R),$$

over the rectangular support $[\underline{r}, \bar{r}]^2$. Notice that the density is indexed by j as a result of the nonidentical densities assumption. This allows the separation rate to deviate from (n-1)/n, the implied separation rate under the i.i.d. assumption.[10]

The turnover rates corresponding to the regional probability masses in Figure 1 have convenient analytical representations. The quit and layoff rates are

$$(6.1) \qquad q(W^o; j) = \int_{W^o}^{\bar{r}} \int_{\underline{r}}^{R} g_j(W, R)\, dW dR$$

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[9] If the n-1 productivity draws are distributed i.i.d., equation (4) reduces to the more familiar form: $\bar{g}(R) = (n-1)g(R)G(R)^{n-2}$.

[10] If the R are distributed multivariate normal, results from Clark (1961) indicate that the distribution of R is approximately normal and the pair (W, R) is distributed approximately bivariate normal. Clark also makes a case for using the normal distribution to approximate the distribution of the maximum of nonnormal random variables.

$$(6.2) \qquad \ell(W^o, R_o; j) = \int_{\underline{r}}^{R_o} \int_{\underline{r}}^{R_o} g_j(W,\ R)\,dWdR \ + \int_{R_o}^{W^o} \int_{\underline{r}}^{R} g_j(W,\ R)\,dWdR.$$

The layoff rate ℓ is the sum of the layoff rates to the nonmarket sector ℓ_o and to other firms ℓ_i:

$$(6.3) \qquad \ell_o(R_o; j) = \int_{\underline{r}}^{R_o} \int_{\underline{r}}^{R_o} g_j(W,\ R)\,dWdR$$

$$(6.4) \qquad \ell_i(W^o, R_o; j) = \int_{R_o}^{W^o} \int_{\underline{r}}^{R} g_j(W,\ R)\,dWdR.$$

The total separation rate s is the mass in the upper half-space plus a triangular region in the lower half-space.

$$(6.5) \qquad s(R_o; j) = \int_{\underline{r}}^{\bar{r}} \int_{\underline{r}}^{R} g_j(W,\ R)\,dWdR \ + \int_{\underline{r}}^{R_o} \int_{R}^{R_o} g_j(W,\ R)\,dWdR.$$

The probability of continued employment is 1-s.

Applying Leibniz's Rule for differentiating through integral expressions, the following comparative statics result:

$$(7.1) \qquad \frac{\partial q}{\partial W^o} = -\frac{\partial \ell}{\partial W^o} = -\frac{\partial \ell_i}{\partial W^o} = -\int_{\underline{r}}^{W^o} g_i(W,\ W^o)\,dW \le 0$$

$$(7.2)\qquad \frac{\partial \ell_o}{\partial W^o} = \frac{\partial s}{\partial W^o} = 0$$

$$(7.3)\qquad \frac{\partial q}{\partial R_o} = 0$$

$$(7.4)\qquad \frac{\partial \ell}{\partial R_o} = \frac{\partial s}{\partial R_o} = \int_{\underline{r}}^{R_o} g_j(R_o,\ R)dR \geq 0$$

$$(7.5)\qquad \frac{\partial \ell_i}{\partial R_o} = -\int_{\underline{r}}^{R} g_j(W,\ R_o)dW \leq 0$$

$$(7.6)\qquad \frac{\partial \ell_o}{\partial R_o} = \int_{\underline{r}}^{R_o} g_j(W,\ R_o)dW + \int_{\underline{r}}^{R_o} g_j(R_o,\ R)dR \geq 0.$$

The nonzero effects are the probabilities of being at the boundaries between the various regions. Shifting W^o and R_o in Figure 1 and noting the change in the various regions should clarify the analytical representations of the partial effects. The higher is a worker's wage, the less (more) likely he is to have a subsequent separation labeled a quit (layoff).[11] But neither the total separation rate nor the layoff rate to the nonmarket sector are

[11]This contrasts with Mortensen's (1978) development of the joint wealth maximizing hypothesis. In his work, the wage serves only to divide the total output between the two parties. He concludes: "Certainly the major empirical implication that distinguishes the joint wealth maximizing hypothesis from the alternative is that both the probability of a worker-initiated separation and the probability of termination by the employer are independent of at least marginal changes in the wage rate" (Mortensen 1978, 583).

influenced by the pre-separation wage W^o. 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.[12]

Indirect Effects

Variables such as employment tenure, experience or age, education, union status, race, and business cycle fluctuations bear no direct consequence in this model. However, these variables are likely to provide important reduced-form or indirect effects through their association with the density function g_j. In the current framework, these variables are modeled as determinants of g_j, W^o, and R_o. In the following cases, it is the differential effect on the variable of interest across g_j, W^o, and R_o that produces the result. I indicate here the most likely of these indirect effects.

First consider level effects which operate primarily on the total separation rate s. Employment tenure, measuring accumulated firm-specific capital, shifts the density g_j and the benchmark wage W^o rightward.[13] This

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[12] It is shown in McLaughlin (1987b, Chapter 6) that the employment rate is decreasing in R_o and is unaffected by W^o.

[13] The introduction of accumulated firm-specific human capital at this stage violates a construct of the model. Turnover behavior is modeled in a sequence of spot markets with decision rules based on only current values of wages and productivities across firms. Such behavior is myopic if there exists accumulated firm-specific human capital. Rational behavior requires (expected) present value calculations over sequences of employment matches. However, the results in this section are not sensitive to this mis-specification; they apply to the properly specified, more complicated model as well.

reduces s in two ways. First, with R_0 constant, there is a lower probability of exiting to the nonmarket sector. Second, the rightward shift pulls the density into the lower half-space. If experience and education move the density northeast (i.e., translate g_j) leaving R_0 unchanged, then the separation rate is falling in these two variables. Like tenure, experience and education reduce s because R_0 is constant, but the second tenure effect--that of improving opportunities with the incumbent employer relative to other market opportunities--does not apply to these variables. So tenure's reduced-form effect on separations should be present even with total market experience controlled for. Analysis of the effect of race (or sex) is similar. If wage differentials by race (sex) are larger in the market sector than in the nonmarket sector, the reduced-form race (sex) effect on the separation rate should indicate less turnover for whites (males).

The second form of indirect effects, termed growth effects, is the result of cyclical and secular growth. The primary influence of growth is on the ratio of quits to layoffs. Productivity growth increases (decreases) the probability that a given separation is to a job paying more (less) than W^o; hence quits are increasing relative to layoffs in general productivity growth.

In McLaughlin (1987b, Chapter 6), I apply the joint wealth maximizing approach to analyze the cyclicality of labor turnover. The business cycle is modeled as a translation of the density g_j: stochastically, opportunities within the incumbent firm increase by exactly the size of the improvement in outside opportunities.[14] Business cycle fluctuations in g_j, as a special case

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[14]This is a mean shift. There is no result with respect to mean preserving spreads.

of general productivity growth, generate the well-known empirical regularities of procyclical quits and countercyclical layoffs.

Such a parameterization is also useful in analyzing secular growth. Workers with steeper wage profiles, or positioned at a steeper point on any given profile should exhibit a higher quit-layoff ratio (McLaughlin 1987b, Chapter 5). If (for reasons outside the model) whites and college graduates have steeper wage profiles, then these two groups are predicted to have higher quit-layoff ratios. Based on concave wage profiles, the quit-layoff ratio is predicted to decrease with experience, and in general with age.[15]

The final indirect effect is intended to capture the difference between union and nonunion workers in turnover behavior. If unions flatten the life-cycle wage profile, and if union workers' best outside opportunities are with other union firms, then the growth effect alone predicts fewer quits and more layoffs of union workers. Alternatively, assume that the only difference between union and nonunion workers is in their abilities to extract rents associated with their matches. Therefore, relax the assumption maintained in the previous section that the worker receives wage offers R equal to his values of productivity r. Each firm i offers the worker some fraction β_i of the match rents. Consider the two firm case. Let the wage payments W and R associated with productivities w and r be generated by the following rent sharing rule:

[15] The growth effects depend critically on the particular benchmark wage. Had I chosen a benchmark which augments W^o for expected growth, the model would not capture the growth-related empirical regularities. See McLaughlin (1987b, Chapter 5).

$$(8.1) \qquad W = R + \beta_w \ (w - R)$$

$$(8.2) \qquad R = W + \beta_r \ (r - W),$$

where β_w and β_r are the two rent-sharing parameters. Treating equations (8.1) and (8.2) as reaction functions, the Nash solution is

$$(8.3) \qquad W = \frac{\beta_w w + \beta_r (1-\beta_w) r}{1 - (1-\beta_w)(1-\beta_r)}$$

$$(8.4) \qquad R = \frac{\beta_w (1-\beta_r) w + \beta_r r}{1 - (1-\beta_w)(1-\beta_r)},$$

with $0 < \beta_i \leq 1$, $i = w, r$. Although this is a simple scheme, it can be shown to have the following desirable properties. First, it induces efficient separations: $W \lesseqgtr R$ as $w \lesseqgtr r$. Second, wage offers are flexible, increasing in own- and other-productivity. Third, a higher share in the optimal match increases the accepted wage.

To analyze the effects of union status within this model, assume the incumbent employer is unionized and the other firm is not unionized; then $\beta_w > \beta_r$. Since the sharing rules (β_w, β_r) are irrelevant for the separation decision, the separation rate is unaffected by union status. However, union status reduces quits relative to layoffs. The union wage premium with the incumbent employer, which is present if and only if the match is optimal, reduces the probability that a given separation is to a higher (than W^o) paying job. Thus union workers are less likely to quit and more likely to be laid off than their nonunion counterparts.[16]

Subsidized-UI Effects

The empirical specifications developed in the remainder of the paper admit other (direct) determinants of labor turnover. The most interesting is the effect of subsidized unemployment insurance (UI). Topel (1983, 1984) presents evidence that subsidized UI increases the incidence of layoffs. His theoretical approach is similar in spirit to mine in that the probability of a layoff is a wealth maximizing decision joint to the firm and worker given imperfect experience rating of employer UI premiums and preferential tax treatment of UI benefits. However, in limiting the analysis to layoffs, Topel cannot investigate the possibility of re-labeling a quit as a layoff in order for the worker to qualify for UI benefits. Subsidized UI generates the incentive to decrease the quit-layoff ratio if some quits exit the market sector. In a more complete specification of my model, R_0 is random and some quits go to the nonmarket sector; hence the ratio of quits to layoffs is predicted to decrease with the degree of subsidization of UI benefits. Of course the incentive to increase the incidence of separations remains. Combining these two, the predicted effects of UI subsidies are to increase the separation and layoff rates, but the effect on quits is ambiguous; the

[16]All three turnover rates are invariant to the magnitude of a common sharing parameter: what drives the result in the text is the change in the incumbent employer's sharing parameter relative to that of the alternative employer. See McLaughlin (1987a) for a more detailed analysis of this rent sharing scheme and its implications for labor turnover.

re-labeling result can reverse an otherwise positive effect.[17] The empirical specification developed and estimated in section 2 of this paper is capable of separately identifying these two components.

A Comparison

Is the joint wealth maximizing specification observationally distinguishable from a rigid wage model? Consider an optimal contract under bilateral asymmetric information which fixes the wage rate in advance (Hall and Lazear 1984). To avoid the costs of renegotiation due to ex post opportunistic behavior, the firm and worker write a binding contract in the initial period governing employment in the subsequent period: if the worker is employed by the firm, he is paid a wage W^1; and either party to the contract can sever the employment relationship, thus separations are unilateral.

In Figure 2, I depict the turnover and employment regions under the fixed-wage contract. There are two differences between Figure 2 and its flexible-wage counterpart, Figure 1. With a fixed wage there are two regions in the lower half-space which generate inefficient separations. Within these two regions, productivity with the incumbent employer exceeds productivity with the best alternative employer; but based on the fixed wage W^1, either the firm lays off the worker or the worker quits the firm. Consequently, separations can be involuntary to the worker or to the firm. This difference is not trivial. Consider for example the case of laid-off workers who buy out the firm to continue their employment. The rigid wage model's property of inefficient separations allows for the success of a worker takeover. The

[17] In more recent work, Topel (1985) includes quits in the analysis. He finds that the incidence of quits is unaffected by UI subsidies.

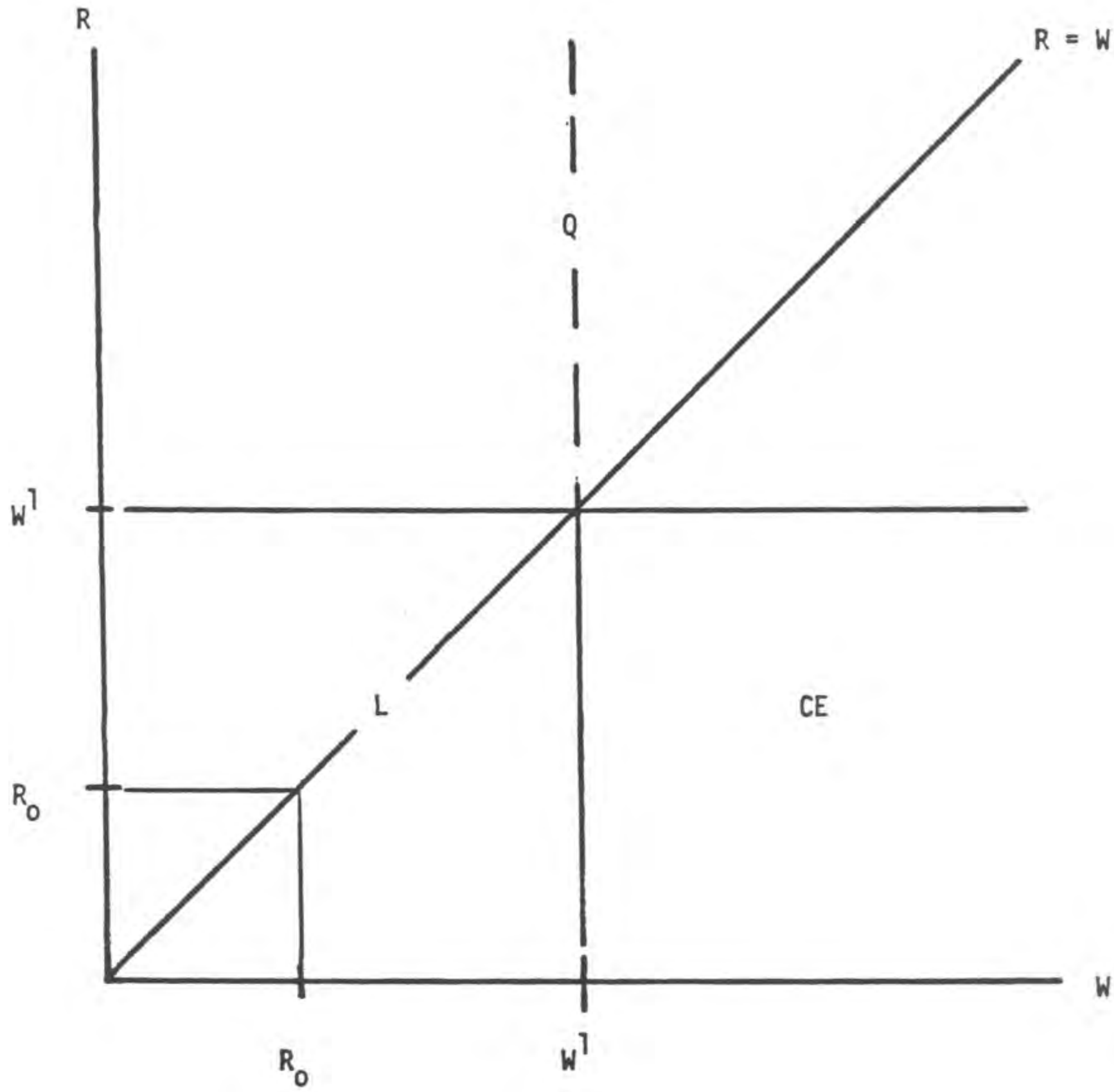

Figure 2. Turnover Regions: Rigid Wage Model

joint wealth maximizing specification is more restrictive in predicting the failure, or indeed absence, of such attempts—not because labor cannot manage itself, but rather--because laid-off workers are more efficient elsewhere.

Assume for the moment that the wage in the rigid wage model is perfectly fixed, so $W^1 = W^0$. An immediate result is that the quit rate is decreasing and the layoff rate is increasing in the wage. However, the effect on the separation rate is distribution specific and in general ambiguous. The joint wealth maximizing specification is more restrictive in generating the invariance of separations to the pre-separation wage.

A second difference between the two models is that W^1, not necessarily W^0, divides the quit and layoff regions in the rigid wage model. In the optimal rigid-wage contract, W^1 need not equal W^0. Rather, the optimal contract conditions W^1 on all symmetrically observable or verifiable information (see Hashimoto and Yu (1980)). Consequently, W^1 is augmented for normal growth in the productivity distribution such as that which occurs over the life cycle: the optimal contract specifies a rigid wage profile not a rigid wage. This subtle difference is quite important. To the extent secular growth, as well as cyclical fluctuations, is forecastable, proxied by observables, or verifiable once revealed by either party to the contract, the first-order effect of growth on the ratio of quits to layoffs is not present in the rigid wage model. Once again, there can be a second-order effect which is distribution specific.

Comparison of my specification of the joint wealth maximizing approach and a simple rigid-wage contract yields two conclusions. On one margin, the joint wealth maximizing specification is more restrictive. On a second margin, the implications of the two models are distinct: the rigid wage model is basically neutral with respect to life cycle productivity growth, and other "general" variation in productivity.

2. Econometric Specification

In this section, I reformulate this joint wealth maximizing model of turnover (i.e., separations and quit-layoff labels) as an empirical model for estimation and testing.[18] The link between the theory in section 1 and the results in section 4 is an econometric specification summarized as follows: The position of the joint density g_j is characterized by a pair of means $(\hat{W}, \hat{R})$, which varies across individuals and through time. Estimates of this pair (or their difference) are used as regressors in two structural probits. First is the separations probit which is governed by the separation decision and its determinants. The second, the labels probit, provides estimates of the determinants of the quit-layoff labels conditional on separation. A direct test of my hypothesis is in the effect of the pre-separation wage W^o and outside opportunities $\hat{R}$ across these two probit equations. The predicted effect of W^o is zero in the separations probit, but the quit rate conditional on separation is falling in W^o; $\hat{R}$ is predicted to increase both separations and the fraction of separations labeled quits.

As in the theoretical section, turnover depends prominently on the triplet (W, R, W^o), and here it also depends on the cost of changing employers, m, and the effect of other variables which influence the quit-layoff distinction, d. Assuming m is proportional to productivity, the

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[18] R_0 is omitted from the empirically implemented model.

multivariate discrete choice model is summarized by the following two index functions:

$$(9.1) \qquad I_1 = \log R - \log W - m$$

$$(9.2) \qquad I_2 = \log R - \log W^{o} - d.$$

The worker is assumed to know the values of the quintuplet (W, R, W^{o}, m, d), but for <u>all observations</u> the analyst observes only W^{o}. However, the other variables are generated as follows:

$$(10.1) \qquad \log W = \log \hat{W} + \epsilon_w = X\beta_w + \epsilon_w$$

$$(10.2) \qquad \log R = \log \hat{R} + \epsilon_R = X\beta_R + \epsilon_R$$

$$(10.3) \qquad m = Z\gamma + \epsilon_m$$

$$(10.4) \qquad d = Y\delta + \epsilon_d.$$

Z is a vector of variables which determine turnover costs or possibly turnover propensities. Y is a vector of other determinants of turnover labels, in particular the degree of subsidization of UI.

Substituting (10.1)-(10.4) into (9.1) and (9.2),

$$(11.1) \qquad I_1 = \log \hat{R} - \log \hat{W} - Z\gamma + \epsilon_1, \quad \epsilon_1 \equiv \epsilon_R - \epsilon_w - \epsilon_m$$

$$(11.2) \qquad I_2 = \log \hat{R} - \log W^{o} - Y\delta + \epsilon_2, \quad \epsilon_2 \equiv \epsilon_R - \epsilon_d.$$

I assume the quadruplet $(\epsilon_R, \epsilon_w, \epsilon_m, \epsilon_d)$ is distributed multivariate normal with zero means and covariance matrix Σ.[19] Therefore, $(\epsilon_1, \epsilon_2) \sim \mathrm{MVN}(0, T'\Sigma T)$, and let $T'\Sigma T \equiv \begin{bmatrix} \sigma_{11} & \sigma_{12} \\ \sigma_{21} & \sigma_{22} \end{bmatrix}$.[20] I also assume that (ϵ_1, ϵ_2) are distributed i.i.d. over time and across workers. In particular, $\log \hat{W}$ and $\log \hat{R}$ must include fixed (and perhaps auto-correlated) person- and match-specific, and common cross-sectional effects.

The worker's turnover behavior is characterized by the following decision rules:

$$(12.1) \qquad S = 1 \text{ if } I_1 > 0, \quad S = 0 \text{ otherwise;}$$

$$(12.2) \qquad Q = 1 \text{ if } I_1 > 0 \ \wedge \ I_2 > 0, \quad Q = 0 \text{ otherwise;}$$

$$(12.3) \qquad L = 1 \text{ if } I_1 > 0 \ \wedge \ I_2 < 0, \quad L = 0 \text{ otherwise.}$$

These correspond to the theoretical model of section 1 (see especially Figure

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[19] While normality is an assumption, it has some merit in this context due to the Clark approximation (Clark 1961).

[20] With $T' \equiv \begin{bmatrix} -1 & 1 & -1 & 0 \\ 0 & 1 & 0 & -1 \end{bmatrix}$, the quadratic form $T'\Sigma T$ is given by the symmetric 2-by-2 matrix

$$\begin{bmatrix} \sigma_{RR} + \sigma_{ww} + \sigma_{mm} - 2\sigma_{Rw} - 2\sigma_{Rm} + 2\sigma_{wm} & \cdot \\ \sigma_{RR} - \sigma_{Rw} - \sigma_{Rm} - \sigma_{Rd} + \sigma_{dw} + \sigma_{dm} & \sigma_{RR} + \sigma_{dd} - 2\sigma_{Rd} \end{bmatrix}.$$

1) modified for turnover costs and the other determinants of the quit-layoff labels. The probit functions associated with equations (12.1)-(12.3) are

$$(13.1) \qquad \Pr[S=0] = \Pr[I_1 < 0] = \Pr[\epsilon_1 < \log \hat{W} - \log \hat{R} + Z\gamma] = \Phi(K_1)$$

$$(13.2) \qquad \begin{aligned} \Pr[Q=1] &= \Pr[I_1 > 0 \ \wedge \ I_2 > 0] \\ &= \Pr[\epsilon_1 > \log \hat{W} - \log \hat{R} + Z\gamma,\ \epsilon_2 > \log W^o - \log \hat{R} + Y\delta] \\ &= 1 - \Phi(K_1) - \Phi(K_2) + \Phi(K_1, K_2) \end{aligned}$$

$$(13.3) \qquad \begin{aligned} \Pr[L=1] &= \Pr[I_1 > 0 \ \wedge \ I_2 < 0] \\ &= \Pr[\epsilon_1 > \log \hat{W} - \log \hat{R} + Z\gamma,\ \epsilon_2 < \log W^o - \log \hat{R} + Y\delta] \\ &= \Phi(K_2) - \Phi(K_1, K_2), \end{aligned}$$

with $K_1 = (\log \hat{W} - \log \hat{R} + Z\gamma)/\sigma_{11}^{1/2}$ and $K_2 = (\log W^o - \log \hat{R} + Y\delta)/\sigma_{22}^{1/2}$. $\Phi(\cdot)$ and $\Phi(\cdot,\cdot)$ are the univariate and bivariate standard normal distribution functions. In Figure 3, the (ϵ_1, ϵ_2) space is decomposed into the three regions defining these probit functions.

On a sample of initially employed individuals, $i = 1, \ldots, I$, the log-likelihood function associated with this model is

$$(14.1) \qquad \begin{aligned} \log \mathcal{L} = &\sum_{i=1}^{I} (1-S_i) \log \Phi(K_1^i) \\ &+ \sum_{i=1}^{I} Q_i \log[1 - \Phi(K_1^i) - \Phi(K_2^i) + \Phi(K_1^i, K_2^i)] \\ &+ \sum_{i=1}^{I} L_i \log[\Phi(K_2^i) - \Phi(K_1^i, K_2^i)]. \end{aligned}$$

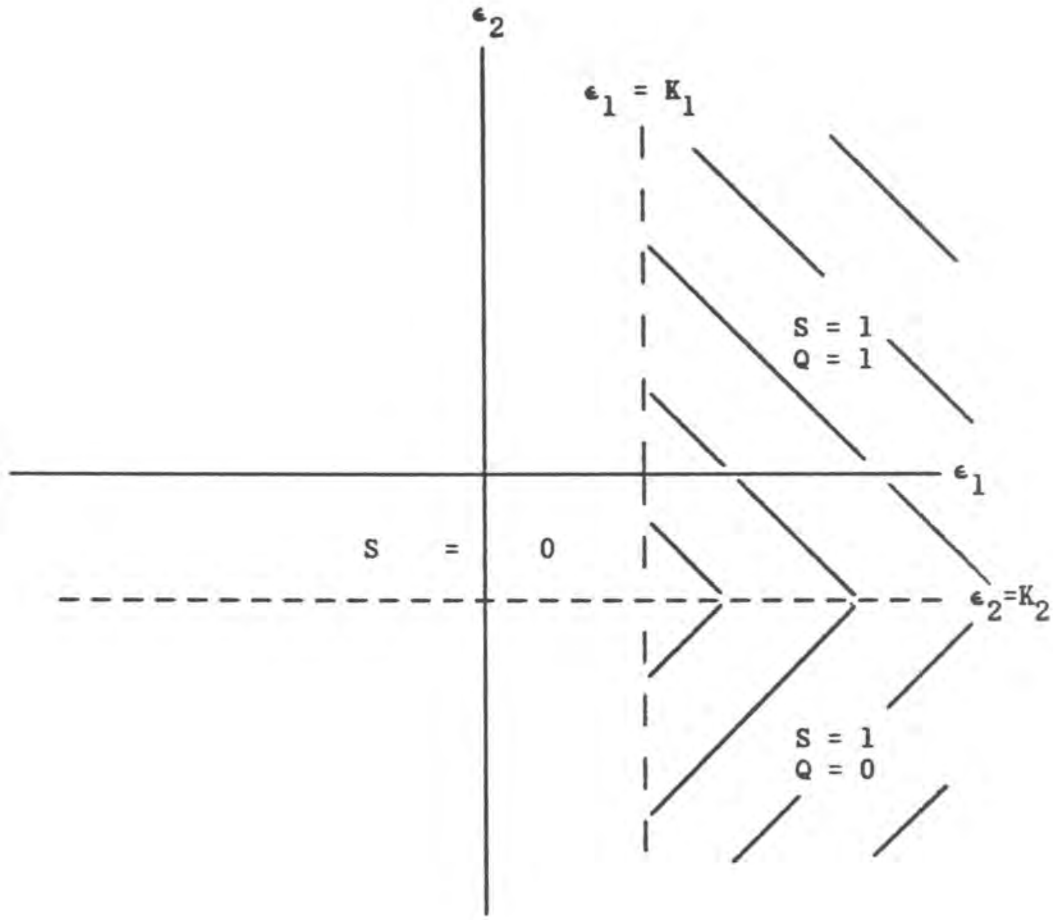

Figure 3. **Bivariate Probit With Selection**

The likelihood function is that of a bivariate probit with selection or one empty cell and could be used directly in producing maximum likelihood estimates. However, by imposing the following restriction, the computational burden is reduced. With $\sigma_{12} = 0$, $\Phi(K_1, K_2) = \Phi(K_1) \cdot \Phi(K_2)$ and (14.1) reduces conveniently to

$$(14.2) \qquad \log \mathcal{L} = \sum_{i=1}^{I} (1-S_i) \log \Phi(K_1^i) + \sum_{i=1}^{I} S_i \log[1 - \Phi(K_1^i)]$$

$$+ \sum_{i=1}^{I} S_i Q_i \log[1 - \Phi(K_2^i)] + \sum_{i=1}^{I} S_i (1-Q_i) \log \Phi(K_2^i).$$

That is, the likelihood function reduces to a probit associated with the separation decision (to be estimated on the full sample) and a labels probit (to be estimated on the subsample of individuals who separate). If the restriction is verified then estimating two distinct probits is full-information maximum likelihood. If the restriction is violated, estimates of the labels probit are biased and inconsistent. The subsample selection criterion is based on S which is endogenous and possibly correlated with the regressors in the labels probit.

Estimation

I employ a modification of Lee's (1978) structural probit method to estimate the model. Estimation of the structural model requires knowledge of the means of the wage-offer distribution ($\log \hat{W}$, $\log \hat{R}$) such that the disturbances (ϵ_W, ϵ_R) exhibit "nice" properties. Of course the pair ($\log \hat{W}$,

$\log \hat{R}$) is not directly available since log W is observable only on the sample with S=0, and log R only on the sample with S=1. However, with consistent estimates of the structural parameters of the wage equations, one can construct the pair ($\log \hat{W}$, $\log \hat{R}$) for each observation in the sample using equations (10.1) and (10.2). The structural parameters of the wage equations are estimated consistently by the two-step selection-bias correction technique (Lee 1978; Heckman 1979). With the pair ($\log \hat{W}$, $\log \hat{R}$) computed, one estimates the structural probit governing the separation decision. In the current context, one also uses $\log \hat{R}$ in the structural probit governing the application of turnover labels.

The structural probit method is modified in two ways. First, I restrict the parameters of the wage equations to be equal across equations: $\boldsymbol{\beta} = \boldsymbol{\beta}_W = \boldsymbol{\beta}_R$. In some contexts it is appropriate to avoid this restriction. For the choices of whether to participate in the union sector (Lee 1978), whether to go to college (Willis and Rosen 1979), or whether to join a white collar occupation (Killingsworth 1985), a sectoral interpretation is appropriate: the rates of return to various characteristics are sector specific. However, the nature of the data employed in the empirical analysis below does not admit a sectoral interpretation. Consider transitions of worker 1 from firm A to firm B and worker 2 from firm B to firm A. Data from firms A and B would be in the wage equations of movers. (Of course, stayers from all firms are pooled into the stayers' wage equation.) Therefore, one cannot attach a sectoral interpretation using these data. Nevertheless, the unrestricted form would be justified if the structural parameters were tenure related. This requires that the return to education, for example, differs between the first year and all other years of employment with the firm. Hence the $\boldsymbol{\beta}_W = \boldsymbol{\beta}_R$ restriction corresponds to an assumption of tenure invariance of the parameters. In principle this restriction is testable.

In considering the pooled wage regression, I allow for two types of regressors: those with values which are invariant to the identity of the employer (X_1)—e.g., education—and those with values which are not employer invariant (X_2)—e.g., employment tenure. The employer-varying characteristics take on values X_w if the worker stays (i.e., if S=0), and X_R if the worker separates (i.e., S=1). Thus $X = (X_1, X_2)$ with $X_2 \equiv (1-S)\cdot X_w + S\cdot X_R$. Letting W' represent the observed wage,

$$(15) \qquad \log W' = (1-S)\cdot \log W + S\cdot \log R$$

$$= X_1\beta_1 + [(1-S)\cdot X_w + S\cdot X_R]\beta_2 + [(1-S)\cdot \epsilon_w + S\cdot \epsilon_R]$$

$$= X_1\beta_1 + X_2\beta_2 + \epsilon.$$

Using standard derivations and definitions from the selection-bias literature (e.g., Heckman 1979),

$$(16.1) \qquad E[\log W' \mid \text{select}] =$$

$$= X_1\beta_1 + X_2\beta_2 + (1-S)\cdot E[\epsilon_w \mid X, S=0] + S\cdot E[\epsilon_R \mid X, S=1]$$

$$= X_1\beta_1 + X_2\beta_2 - \alpha_0[(1-S)\cdot \lambda_0(-K_1)] + \alpha_1[S\cdot \lambda_1(-K_1)]$$

$$(16.2) \qquad \lambda_0(-K_1) \equiv - \frac{\phi(-K_1)}{1 - \Phi(-K_1)}, \qquad \lambda_1(-K_1) \equiv \frac{\phi(-K_1)}{\Phi(-K_1)},$$

$$\alpha_0 \equiv \frac{\sigma_{ww} - \sigma_{wR} - \sigma_{wm}}{\sigma_{11}^{1/2}}, \qquad \alpha_1 \equiv \frac{\sigma_{RR} - \sigma_{wR} - \sigma_{mR}}{\sigma_{11}^{1/2}}.$$

Assuming $\sigma_{ww} = \sigma_{RR}$ and $\sigma_{wm} = \sigma_{mR}$ (based on the data), and defining $\lambda(-K_1) \equiv -(1-S)\cdot\lambda_0(-K_1) + S\cdot\lambda_1(-K_1)$, the following pooled wage regression obtains:

$$\text{(17)} \qquad \log W' = X_1\beta_1 + X_2\beta_2 + \alpha\lambda(-K_1) + \nu$$

where ν is mean zero and uncorrelated with the vector (X, λ). Consequently, the pooled wage regression is estimable by Heckman's two-step procedure <u>on all observations</u> using $\lambda(-K_1)$ as the selectivity variable.[21]

There is a second procedure which can be used to estimate the structural parameters β. It is based on a simultaneous equations—as opposed to an omitted variables—interpretation of the estimation problem. Note that the variables in the vector $X_2 \equiv (1-S)\cdot X_w + S\cdot X_R$ are in general correlated with $\epsilon \equiv (1-S)\cdot\epsilon_w + S\cdot\epsilon_R$, because S is endogenous and depends on ϵ_w and ϵ_R. The instrumental variables (IV) estimator employed in the union-nonunion context by Robinson and Tomes (1984) and Duncan and Leigh (1985) accounts for the simultaneity. The estimator uses $\hat{X}_2 \equiv (1-\hat{S})\cdot X_w + \hat{S}\cdot X_R$ as the vector of instrumental variables for X_2, where $\hat{S}$ is the prediction from a reduced-form probability model.

Estimates from both procedures are reported in section 4 below. Both

[21]Since $\alpha \equiv (\sigma_{ww} - \sigma_{wR} + \sigma_{wm})/\sigma_{11}^{1/2} \equiv (\sigma_{RR} - \sigma_{wR} + \sigma_{mR})/\sigma_{11}^{1/2}$, α must be nonnegative. $\alpha < 0$ would imply that the variance of ϵ_1 is negative.

sets of estimates are quite similar to the OLS estimates.

3. Data

To estimate and test the turnover model, I employ a sample from the Panel Study of Income Dynamics (PSID) spanning the years 1975-1980. The six years of data are pooled into five yearly transition subsamples: 1975-1976, 1976-1977, ..., 1979-1980. For each of the five subsamples, I refer to the first year as the initial period and the second year as the subsequent period. The sample consists of 10,922 initially employed, male, household heads, age 18 to 58. Of the 10,922 observations, 10,469 are employed in the subsequent period, and 16 percent of the sample separate from their employers. Details of the sample exclusions and treatment of missing observations are provided in McLaughlin (1987b, Appendix 1).[22]

The following variables are used in the empirical analysis. For the turnover variables, the PSID asks each respondent who has changed jobs in the intervening year the reason for his job mobility. I combine the responses "laid off; fired" and "company folded" to define the layoff variable. Quits include resignations, separations to retire, and responses such as "just wanted to change jobs." The real wage variable (for both periods) is the log of hourly pay on the main job deflated to 1972 dollars by the GNP deflator. The extra variables (Z) included in the separations probit are age, dummies for industry, occupation, union membership (or covered by a union contract) and marital status, and a measure of business cycle fluctuations from

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[22] A discussion of how X_W is computed for movers and how X_R is imputed for stayers is also included in the referenced appendix.

McLaughlin (1987b, Chapter 6). The cyclical variable is computed as the four-quarter average of deviations of (log) real GNP from a cubic time trend. The additional variables (Y) in the labels probit are limited to the cyclical variable at this stage, but could include a variable measuring the degree of subsidization of unemployment insurance benefits.

The variables in the pooled wage regression are: education, a dummy variable to mitigate the effects of censoring the education variable at 17, experience (years employed since age 18), tenure (years with the current employer), dummy variables for industry, race, region, marital status, government employment, and union status, and the cyclical variable. Employment tenure, and the industry, region, government employment, and union dummy variables are the variables in X_2; all other wage equation variables are in X_1. All wage equation variables take on values from the subsequent period.

Summary statistics for each subsample and the pooled sample are reported in Table 1.

4. Empirical Results

The empirical results to be discussed in this section are displayed in Tables 2-4. These results include OLS, Heckman Two-Step, and IV estimates of the wage equation, and estimates of the structural turnover model by two structural probit equations and by maximum likelihood.

Various estimates of the wage equation are reported in Table 2. Aside from the estimated effect of employment tenure in the Two-Step column, the estimates exhibit standard signs and magnitudes. For example, the rate of return to education is about 6 percent; the rates of return to labor market experience and employment tenure (at the sample means of 15 and 8) are,

TABLE 1

SUMMARY STATISTICS
PSID, 1975–1980

VARIABLE	1975–1976		1976–1977		1977–1978		1978–1979		1979–1980		POOLED	
	0	1	0	1	0	1	0	1	0	1	0	1
Turnover:												
S	–	0.15 (0.35)	–	0.15 (0.36)	–	0.15 (0.36)	–	0.16 (0.37)	–	0.18 (0.39)	–	0.16 (0.37)
Q[a]	–	0.50 (0.50)	–	0.58 (0.50)	–	0.66 (0.47)	–	0.69 (0.46)	–	0.63 (0.48)	–	0.63 (0.48)
L[a]	–	0.50 (0.50)	–	0.42 (0.50)	–	0.34 (0.47)	–	0.31 (0.46)	–	0.37 (0.48)	–	0.37 (0.48)
Employment:												
E	1	0.95 (0.21)	1	0.96 (0.20)	1	0.96 (0.20)	1	0.97 (0.18)	1	0.95 (0.22)	1	0.96 (0.20)
U	0	0.03 (0.17)	0	0.03 (0.16)	0	0.03 (0.16)	0	0.02 (0.14)	0	0.03 (0.18)	0	0.03 (0.16)
WAGES[b]	3.72 (1.39)	3.93 (1.44)	4.19 (1.68)	4.41 (1.63)	4.24 (1.62)	4.65 (2.37)	4.50 (2.39)	4.68 (2.45)	4.50 (2.43)	4.66 (2.49)	4.28 (2.03)	4.52 (2.20)
EDUCAT	10.98 (2.91)	10.98 (2.90)	11.98 (3.09)	11.98 (3.08)	11.98 (2.99)	11.99 (2.99)	12.10 (2.88)	12.10 (2.89)	12.20 (2.77)	12.21 (2.76)	11.93 (2.95)	11.93 (2.94)
EXPER	16.35 (11.92)	17.03 (11.02)	16.21 (10.98)	16.21 (10.97)	15.16 (10.71)	15.17 (10.71)	14.20 (10.45)	14.22 (10.44)	13.06 (10.03)	13.02 (10.00)	14.81 (10.78)	14.89 (10.67)
TENURE[b]	6.53 (7.08)	8.14 (7.95)	7.67 (8.03)	8.55 (8.52)	7.70 (8.26)	8.51 (7.88)	8.01 (6.76)	8.68 (7.04)	7.72 (6.84)	8.55 (7.25)	7.62 (7.43)	8.52 (7.69)
GOVT[b]	0.19 (0.60)	0.17 (0.37)	0.22 (0.41)	0.21 (0.40)	0.20 (0.40)	0.21 (0.40)	0.21 (0.41)	0.21 (0.40)	0.20 (0.40)	0.21 (0.40)	0.20 (0.43)	0.21 (0.40)
UNIONS	0.44 (0.50)	0.47 (0.50)	0.36 (0.48)	0.36 (0.50)	0.36 (0.48)	0.37 (0.48)	0.37 (0.48)	0.36 (0.48)	0.36 (0.48)	0.35 (0.48)	0.37 (0.48)	0.37 (0.48)

Industry (I):[b]

1. Agriculture	0.03 (0.17)	0.03 (0.18)	0.03 (0.16)	0.03 (0.16)	0.03 (0.16)	0.02 (0.15)	0.02 (0.15)	0.02 (0.15)	0.02 (0.15)	0.02 (0.15)	0.03 (0.16)	0.03 (0.16)
2. Mining	0.01 (0.12)	0.02 (0.13)	0.01 (0.10)	0.01 (0.12)	0.01 (0.12)	0.01 (0.11)	0.02 (0.12)	0.02 (0.12)	0.02 (0.12)	0.02 (0.12)	0.01 (0.12)	0.01 (0.12)
3. Manufacturing	0.41 (0.49)	0.40 (0.49)	0.35 (0.47)	0.34 (0.47)	0.35 (0.48)	0.37 (0.48)	0.36 (0.48)	0.36 (0.47)	0.36 (0.47)	0.36 (0.48)	0.36 (0.48)	0.36 (0.48)
4. Construction	0.12 (0.32)	0.10 (0.30)	0.08 (0.27)	0.09 (0.30)	0.09 (0.28)	0.09 (0.28)	0.09 (0.29)	0.10 (0.29)	0.10 (0.30)	0.09 (0.28)	0.09 (0.29)	0.09 (0.29)
5. Transp.-Comm.	0.12 (0.33)	0.13 (0.34)	0.10 (0.30)	0.10 (0.30)	0.10 (0.30)	0.11 (0.31)	0.10 (0.30)	0.11 (0.31)	0.11 (0.31)	0.11 (0.32)	0.11 (0.31)	0.11 (0.31)
6. Trade	0.10 (0.30)	0.10 (0.30)	0.13 (0.33)	0.13 (0.33)	0.13 (0.33)	0.12 (0.32)	0.12 (0.32)	0.11 (0.31)	0.12 (0.32)	0.12 (0.32)	0.12 (0.32)	0.12 (0.32)
7. Finance, etc.	0.01 (0.12)	0.02 (0.13)	0.03 (0.17)	0.03 (0.17)	0.03 (0.16)	0.03 (0.17)	0.03 (0.16)	0.03 (0.17)	0.03 (0.17)	0.03 (0.17)	0.03 (0.16)	0.03 (0.16)
8. Repair Service	0.05 (0.21)	0.03 (0.18)	0.04 (0.18)	0.03 (0.18)	0.03 (0.18)	0.03 (0.18)	0.04 (0.16)	0.03 (0.18)	0.03 (0.18)	0.03 (0.16)	0.03 (0.18)	0.03 (0.18)
9. Personal Serv.	0.01 (0.10)	0.01 (0.10)	0.01 (0.11)	0.01 (0.10)	0.01 (0.10)	0.01 (0.09)	0.01 (0.09)	0.01 (0.07)	0.01 (0.07)	0.01 (0.09)	0.01 (0.09)	0.01 (0.09)
10. Entertainment	0.00 (0.06)	0.01 (0.09)	0.01 (0.09)	0.01 (0.08)	0.00 (0.07)	0.00 (0.07)	0.01 (0.07)	0.00 (0.07)	0.00 (0.07)	0.01 (0.07)	0.01 (0.07)	0.01 (0.07)
11. Professional Service	0.05 (0.23)	0.06 (0.23)	0.11 (0.31)	0.11 (0.31)	0.11 (0.31)	0.11 (0.31)	0.11 (0.31)	0.11 (0.31)	0.11 (0.31)	0.11 (0.32)	0.10 (0.30)	0.10 (0.30)
12. Gov't, etc.	0.08 (0.26)	0.08 (0.28)	0.11 (0.31)	0.11 (0.32)	0.11 (0.31)	0.10 (0.30)	0.10 (0.30)	0.10 (0.30)	0.10 (0.30)	0.09 (0.29)	0.10 (0.30)	0.10 (0.30)
Occupation (O):[b]												
1. Professional	0.07 (0.25)	0.06 (0.24)	0.14 (0.35)	0.14 (0.35)	0.13 (0.34)	0.14 (0.35)	0.15 (0.35)	0.15 (0.35)	0.14 (0.35)	0.15 (0.35)	0.13 (0.34)	0.13 (0.34)
2. Managers (NSE)	0.03 (0.16)	0.04 (0.19)	0.11 (0.31)	0.12 (0.32)	0.11 (0.32)	0.11 (0.31)	0.11 (0.31)	0.12 (0.32)	0.11 (0.31)	0.12 (0.33)	0.10 (0.30)	0.11 (0.31)
3. Managers (SE)	0.00 (0.00)	0.00 (0.07)	0.00 (0.00)	0.01 (0.09)	0.00 (0.04)	0.01 (0.08)	0.00 (0.02)	0.01 (0.07)	0.00 (0.03)	0.00 (0.05)	0.00 (0.02)	0.01 (0.07)

TABLE 1 --Continued

VARIABLE	1975-1976		1976-1977		1977-1978		1978-1979		1979-1980		POOLED	
	0	1	0	1	0	1	0	1	0	1	0	1
4. Clerical	0.10 (0.30)	0.11 (0.31)	0.11 (0.31)	0.10 (0.30)	0.09 (0.29)	0.10 (0.29)	0.09 (0.29)	0.08 (0.28)	0.08 (0.28)	0.09 (0.28)	0.09 (0.29)	0.09 (0.29)
5. Craftsmen	0.28 (0.45)	0.29 (0.46)	0.24 (0.43)	0.26 (0.44)	0.26 (0.44)	0.27 (0.44)	0.26 (0.44)	0.27 (0.44)	0.27 (0.44)	0.25 (0.43)	0.26 (0.44)	0.27 (0.44)
6. Operatives	0.30 (0.46)	0.31 (0.46)	0.23 (0.42)	0.21 (0.41)	0.23 (0.42)	0.23 (0.42)	0.24 (0.43)	0.23 (0.42)	0.24 (0.42)	0.24 (0.43)	0.24 (0.43)	0.24 (0.43)
7. Laborers	0.19 (0.39)	0.17 (0.38)	0.15 (0.35)	0.15 (0.35)	0.15 (0.36)	0.13 (0.33)	0.44 (0.35)	0.13 (0.34)	0.15 (0.35)	0.13 (0.34)	0.15 (0.36)	0.14 (0.35)
8. Farmers	0.00 (0.05)	0.00 (0.06)	0.00 (0.05)	0.00 (0.04)	0.00 (0.03)	0.00 (0.06)	0.00 (0.04)	0.00 (0.04)	0.00 (0.00)	0.00 (0.02)	0.00 (0.04)	0.00 (0.04)
9. Military	0.03 (0.17)	0.00 (0.07)	0.02 (0.13)	0.02 (0.13)	0.02 (0.13)	0.02 (0.13)	0.01 (0.12)	0.02 (0.12)	0.01 (0.11)	0.01 (0.11)	0.02 (0.13)	0.01 (0.12)
AGE	34.81 (11.15)	35.84 (11.14)	35.05 (11.03)	36.02 (11.04)	34.90 (10.89)	35.81 (10.90)	34.65 (10.69)	35.67 (10.70)	34.26 (10.45)	35.29 (10.45)	34.71 (10.80)	35.70 (10.81)
RACE	0.59 (0.49)		0.66 (0.47)		0.65 (0.48)		0.65 (0.48)		0.64 (0.48)		0.46 (0.48)	
MARRIED	0.87 (0.34)	0.88 (0.33)	0.87 (0.34)	0.87 (0.34)	0.85 (0.35)	0.86 (0.35)	0.86 (0.35)	0.86 (0.35)	0.84 (0.37)	0.83 (0.38)	0.86 (0.35)	0.86 (0.35)
SOUTH	0.47 (0.50)	0.47 (0.50)	0.43 (0.50)	0.43 (0.50)	0.43 (0.50)	0.43 (0.49)	0.43 (0.50)	0.44 (0.50)	0.44 (0.50)	0.44 (0.50)	0.44 (0.50)	0.44 (0.50)
CYCLIC	-	-3.31 (0.00)	-	-1.57 (0.00)	-	1.07 (0.00)	-	2.43 (0.00)	-	0.92 (0.00)	-	0.25 (1.90)
N N-EMPLOYED	1417	1417 1353	2158	2158 2072	2298	2298 2205	2456	2456 2377	2593	2593 2462	10,922	10,922 10,469

[a]Q and L are defined on the subsample of movers (S=1).

[b]WAGES, TENURE, GOVT, and industry and occupation dummy variables are defined on the employed subsample (E=1).

respectively, about 0.7 percent and from 1.1 to 1.7 percent, so a year of continued employment increases the wage from 1.9 to 2.3 percent. However, the two-step estimate of the effect of employment tenure is small and is not statistically significant.

The wage equation estimates reported in Table 2 exhibit several important features. That "corrected" estimates exhibit familiar signs and magnitudes counters several recent contributions to the literature on estimating wage equations in the presence of endogenous turnover. Altonji and Shakotko (1987), Abraham and Farber (1987), and Marshall and Zarkin (1987) use various techniques to control for endogenous turnover, and find a negligible structural effect of employment tenure. Closest to my approach is Marshall and Zarkin's; however, their estimates differ markedly from mine. Using NLS data, Marshall and Zarkin correct for selection bias due to systematic turnover behavior in estimating the wage equations of movers and stayers (or new workers and more senior workers). Their point estimates indicate that the wage-tenure profile is a <u>negatively-sloped</u> convex function, but the coefficients are not statistically significant. Two features account for the difference between our estimates. First, the selection-bias corrected estimates reported in the two-step column of Table 2 indicate a small, statistically insignificant effect of employment tenure on the wage rate. But this is not the principal difference: Marshall and Zarkin do not restrict β_W to equal β_R. My (unreported) unrestricted estimates using either the two-step or IV procedure on the PSID data indicate <u>significant, negative</u> returns to

TABLE 2

WAGE EQUATION ESTIMATES[a]

Regressor	OLS	Two Step[b]	IV1[c]	IV2[d]
constant	0.294 (0.023)	0.412 (0.025)	0.235 (0.023)	0.310 (0.024)
EDUCAT	0.061 (0.001)	0.060 (0.001)	0.064 (0.002)	0.060 (0.002)
ED17	0.163 (0.016)	0.155 (0.015)	0.152 (0.016)	0.156 (0.016)
EXPER	0.018 (0.001)	0.020 (0.001)	0.014 (0.001)	0.013 (0.001)
$EXPER^2/100$	-0.033 (0.003)	-0.035 (0.003)	-0.023 (0.004)	-0.022 (0.004)
TENURE	0.014 (0.001)	0.003 (0.002)	0.023 (0.002)	0.026 (0.002)
$TENURE^2/100$	-0.021 (0.004)	0.002 (0.004)	-0.052 (0.005)	-0.059 (0.005)
SOUTH	-0.124 (0.007)	-0.126 (0.007)	-0.125 (0.008)	-0.127 (0.008)
GOVT	-0.108 (0.012)	-0.117 (0.012)	-0.122 (0.013)	-0.113 (0.013)
UNION	0.128 (0.007)	0.121 (0.007)	0.142 (0.008)	0.121 (0.008)
RACE	0.124 (0.008)	0.130 (0.008)	0.127 (0.008)	0.127 (0.008)
MARRIED	0.084 (0.010)	0.076 (0.010)	0.078 (0.010)	0.073 (0.010)
CYCLIC	0.013 (0.002)	0.016 (0.002)	0.012 (0.002)	0.011 (0.002)
Industry[e]	yes [75.44]	yes [76.27]	yes [51.51]	yes [71.45]
LAMBDA		-0.135 (0.011)		
R^2	.449	.457	.437	.448

[a]10,043 observations per regression; the 879 observations with missing wage data are deleted. Standard errors are in parentheses.

[b]OLS standard errors.

[c]Instrumental variables are computed using estimates from the reduced-form probit governing the separation decision. OLS standard errors.

[d]Instrumental variables are computed using estimates from the reduced-form linear probability model governing the separation decision. IV standard errors.

[e]In brackets are the F-statistics for tests of the joint significance of the eleven industry dummy variables.

employment tenure.[23] Indeed, the restriction $\beta_W = \beta_R$ can be rejected. Nevertheless, I do not relax this restriction for two reasons. First, the unrestricted estimates are not sufficiently robust, since the theoretical case for tenure-varying parameters is not strong.[24] Second, estimates of log $\hat{W}$ and log $\hat{R}$ from these unrestricted wage equations embody an unfortunate property: log $\hat{R}$ is high relative to log $\hat{W}$ for high tenure, low separation rate workers. With this property the baseline model of efficient turnover cannot fit.

The parameter estimates from the IV2 column of Table 2 are used to compute log $\hat{W}$ and log $\hat{R}$ by equations (10.1) - (10.2) for all 10,922 observations. For the full sample, the means (standard deviations) of the two variables are 1.41 (0.30) and 1.29 (0.25), respectively. These statistics indicate that the wage-offer distribution, on average, is positioned below the R=W ray in the (W, R) plane (see Figure 1). The difference log $\hat{W}$ - log $\hat{R}$ is on average twice as large for stayers as for separations; workers whose opportunities are on average relatively better with the incumbent employer are less likely to separate. The correlation between log $\hat{W}$ and log $\hat{R}$ is .89 on the full sample, but somewhat lower on the sample of separations. That the correlations are quite high is consistent with the presence of strong

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[23]Problems with the estimates of the unrestricted wage equation of stayers also include: an unusually small return to schooling, negative return to experience, and negative effects of both union and marital status. The problems in the unrestricted wage equation of movers are less severe.

[24]This suggests that the underlying model is not properly specified. With a quarter century of research supporting rising wage-experience and wage-tenure profiles, and without a theoretical case for tenure-varying parameters, it might well be better to impose the restriction rather than to rely on the tale of the data. This topic certainly deserves further investigation.

general-human- capital components which shift the wage-offer distribution.[25]

Probit estimates of the two equation (separations and labels), structural turnover model are reported in columns (1) and (3) of Table 3. (The estimates in columns (2) and (4) are included for comparison.) First consider the separations probit.[26] The parameter estimates indicate that the separation rate is falling in the difference between opportunities within and outside the incumbent firm, $\log \hat{W} - \log \hat{R}$. This evidence supports the model of separations. First, the theory implies that $\log \hat{W}$ and $\log \hat{R}$ enter as a difference. The likelihood ratio test of this restriction does not reject the null hypothesis. Second, the separation rate s is quite responsive to this difference. For instance, increasing $\hat{W}$ by 10 percent, holding $\hat{R}$ fixed, reduces the separation rate 2.5 percentage points. However, the probit estimates in column (1) do not support the underlying model of efficient turnover. Counter to the model's implication, the separation rate is falling in the pre-separation wage $\log W^{o}$. The probit coefficient -0.367 implies that a 10 percent increase in the pre-separation wage, holding the wage-offer distribution fixed, reduces the separation rate 0.9 percentage points. Although this effect is sizeable, it is only one-third the magnitude of the effects of $\log \hat{W}$ and $\log \hat{R}$. The robustness of the finding to the presence of match-specific capital is discussed below.

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[25] This evidence does not refer to the shape of the iso-probability contour in Figure 1. Rather it refers to the position of the contour.

[26] In the text, I report each variable's partial effect which is derived by multiplying the estimated probit coefficient by a turnover-rate-specific factor of proportionality—the value of the probability density function evaluated at the mean of the turnover rate. The factor of proportionality is .24913 for the separations probit, and .37748 for the labels probit.

TABLE 3

PROBIT ESTIMATES OF THE TURNOVER MODEL[a]

Regressor[b]	Separations (S=1)[c] (1)	Separations (S=1)[c] (2)	Labels (Q=1)[d] (3)	Labels (Q=1)[d] (4)
constant	0.411 (0.108)	0.567 (0.080)	−0.489 (0.161)	0.124 (0.091)
log $\hat{W}$	−1.014 (0.138)		0.291 (0.207)	
log $\hat{R}$	1.084 (0.149)		0.420 (0.209)	
log W^o	−0.367 (0.049)	−0.394 (0.044)	−0.089 (0.093)	0.148 (0.073)
AGE	−0.021 (0.002)	−0.025 (0.002)		
UNION	−0.132 (0.037)	−0.183 (0.036)		
MARRIED	−0.219 (0.041)	−0.201 (0.040)		
CYCLIC	0.028 (0.008)	0.023 (0.008)	0.074 (0.016)	0.079 (0.016)
Industry[e]	yes [117.0]	yes [101.8]	no	no
Occupation[e]	yes [19.8]	yes [20.2]	no	no
log-likelihood	−4,341.9	−4,370.7	−1,138.8	−1,149.5
psuedo R^2	.100	.094	.021	.012

[a]Asymptotic standard errors in parentheses. Psuedo R^2 is McFadden's measure.

[b]AGE, UNION, MARRIED, and industry and occupation dummy variables take on values from the initial period; that is, their pre-separation values.

[c]10,922 observations per probit regression.

[d]1,760 observations per probit regression.

[e]In brackets are the likelihood ratio statistics for tests of the joint significance of either the eleven industry or eight occupation dummy variables.

Several other variables are included in the separations probit. To the extent log $\hat{W}$ and log $\hat{R}$ control for the distribution of wage offers, the probit coefficients associated with these other variables are estimates of direct effects. The separation rate declines with age by half a percentage point per year. Union status reduces the separation rate by approximately 3 percentage points. Consistent with the idea of family decision making limiting "migration" (Mincer 1978), marriage reduces the separation rate by 5 percentage points. There is also evidence of a procyclical separation rate; a cyclical expansion of 1 percent increases the separation rate by 0.7 percentage points. Consequently, the probit estimates indicate that age, union and marital status, and cyclical fluctuations influence the separation rate directly.

Turn now to the estimates of the labels probit—the probability that a given separation is a quit--reported in column (3) of Table 3. The theory implies that, with the distribution of wage offers and the pre-separation wage controlled for, the only other variable to effect quits relative to layoffs is the degree to which unemployment insurance benefits are subsidized. With this variable not yet incorporated into the empirical analyis, I limit the extra variables in the labels probit to the cyclical variable as the cyclicality of quits and layoffs is known to be strong.

The parameter estimates reported in column (3) of Table 3 offer only weak support for the joint wealth maximizing hypothesis. Consistent with the model, there is evidence of a sizeable positive effect of outside opportunities log $\hat{R}$ on quits conditional on separation: a 10 percent increase in $\hat{R}$ increases the quit rate conditional on separation by 1.5 percentage points. The effect of log $\hat{W}$, which is predicted to be zero, is positive but is not statistically significant. Hence both of these parameter estimates support the model. However, the effect of the pre-separation wage W^o is surprisingly weak: a 10 percent increase in the pre-separation wage, holding the wage-offer distribution fixed, reduces the quit rate conditional on separation by only 0.3 percentage points. Furthermore, the probit coefficient is not statistically significant. This appears, at least in part, to be an artifact of the particular data set. The results in McLaughlin (1987b, Chapter 4) indicate that, unlike other data sets, quits and layoffs are falling in the pre-separation wage <u>by the same magnitude</u>. That a cyclical expansion of 1 percent increases the quit rate conditional on separation 2.8 percentage points is evidence of a cyclical effect which is not accounted for by cyclical wage variation.

I also estimate the unrestricted model, the bivariate probit with selection, by maximum likelihood. The maximum likelihood estimate of the correlation "across probits" is 0.003 (0.1328) which validates the $\sigma_{12}=0$ restriction adopted at the outset. In any case, the parameter estimates remain essentially unchanged and are not reported.

Estimates of the structural turnover model offer limited support for the joint wealth maximizing hypothesis. Although the coefficients on log $\hat{W}$ and log $\hat{R}$ in the separations probit are as expected, the principal implications regarding the quit-layoff distinction refer to log W^o for separations and both log W^o and log $\hat{R}$ for labels. Although the magnitude of log W^o's effect on the quit rate conditional on separation is small, it is of the predicted sign; and log $\hat{R}$'s effect in the labels probit supports the model. Hence a rejection of the joint wealth maximizing hypothesis must be driven from the estimated effect of $\log W^o$ on the separation rate. However this result is not robust to the presence of unobserved, fixed, firm-specific human capital.

Fixed, Firm-Specific Capital

The test of the null hypothesis of a zero coefficient on $\log W^o$ in the separations probit is not robust to the presence of fixed (as opposed to accumulated), firm-specific human capital. The unobserved presence of such match capital, which unlike accumulated firm-specific capital is not picked up by $\log \hat{W}$ and $\log \hat{R}$, biases the coefficient on the pre-separation wage. Since match capital lowers the separation rate and is positively correlated with the pre-separation wage, the estimated effect of $\log \hat{W}$ is biased downward.[27] This is a problem with the predictions from the wage equation: $\log \hat{W}$ and $\log \hat{R}$. If the predictions include match-capital components, the omitted variable problem vanishes.

A thorough analysis incorporating match-specific components is beyond the scope of the current work, but a simple extension should help gauge the severity of the problem. In particular, I limit the modification to match capital with the incumbent employer since it is from here that the problem arises. Letting μ_{wi} denote the match-specific effect, and assuming time- and tenure-invariant parameters, the wage equation becomes

$$(18) \qquad \log W_i^t = X_i^t\beta + \mu_{wi} + \epsilon_{wi}^t, \qquad t = 0, 1, \quad i = 1, \ldots, I.$$

[27] Fixed general human capital—a fixed person effect in the wage equation—is not problematic because it does not effect the separation rate.

Thus

$$(19)\qquad \log W_i^1 = X_i^1\beta + \mu_{wi} + \epsilon_{wi}^1$$

$$= X_i^0\beta + \mu_{wi} + \epsilon_{wi}^0 + (X_i^1 - X_i^0)\beta + (\epsilon_{wi}^1 - \epsilon_{wi}^0)$$

$$= \log W_i^0 + \Delta X_i\beta + \Delta\epsilon_i = \log \hat{W}_i + \Delta\epsilon_{wi}, \quad i = 1, \ldots, I,$$

with $\Delta X_i \equiv X_i^1 - X_i^0$, and $\Delta\epsilon_{wi} \equiv \epsilon_{wi}^1 - \epsilon_{wi}^0$. By using $\log W_i^0$ in the computation, the match-specific effect is included in $\log \hat{W}_i$.

A problem in the construction of $\log \hat{W}$ might remain. Since $\log \hat{W} = \log W^0 + \Delta X\beta$, it includes ϵ_w^0. This is inappropriate if ϵ_w^t is distributed i.i.d. over time. However, if ϵ_w^t follows a random walk, this is the appropriate computation.

Probit estimates of the structural turnover model using the modified control for the incumbent employer's wage-offer distribution are reported in Table 4.[28] Since one cannot reject the hypothesis of $\log \hat{W}$ and $\log \hat{R}$ entering as a difference, I report the restricted estimates. Consistent with the presence of fixed, firm-specific human capital, the effect of the pre-separation wage on the separation rate drops by one-third using the modified $\log \hat{W}$ variable, but remains statistically significant. The new

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[28] One could estimate β by regressing wage growth on the differenced regressors and use the estimates to construct $\log \hat{W}$. Alternatively, I use the IV2 estimates from Table 2 to compute $\log \hat{W}$ directly.

TABLE 4

PROBIT ESTIMATES OF THE TURNOVER MODEL[a]
--Modified for Match Capital--

Regressor[b]	Separations (S=1)[c]	Labels (Q=1)[d]
constant	0.435 (0.101)	−0.446 (0.159)
log $\hat{W}$	−0.172 (0.080)	0.196 (0.811)
log $\hat{R}$	0.172 (0.080)	0.629 (0.145)
log W^o	−0.259 (0.076)	−0.228 (0.804)
AGE	−0.025 (0.002)	
UNION	−0.194 (0.037)	
MARRIED	−0.216 (0.041)	
CYCLIC	0.021 (0.008)	0.077 (0.018)
Industry	yes	no
Occupation	yes	no
log-likelihood	−4,368.3	−1,139.8
psuedo R^2	.094	.021

NOTES: See Table 3.

estimates of the labels probit are somewhat stronger. The coefficient on log $\hat{W}$, which the model predicts to be zero, drops from .291 to .196 and remains statistically insignificant. The predicted positive effect of log $\hat{R}$ becomes even stronger, and the predicted negative effect of the pre-separation wage strengthens from -.089 to -.228 but remains insignificant.

These modified results offer limited support for, but formally reject, the joint wealth maximizing hypothesis. The structural probit estimates of the separation decision indicate that estimates of several parameters are not robust to the presence of match capital, and the power of the tests is low. Since the model does quite well in accounting for the empirical regularities of the quit-layoff distinction (McLaughlin 1987b, Chapters 4-6), a more thorough incorporation of the effect of match capital is justified.

5. Summary and Conclusion

The approach I have taken in determining whether quits and layoffs are economically distinct labor force transitions is to develop a strong case for "no meaningful distinction" and to determine whether such a model's refutable implications are consistent with the evidence. The theoretical model appends turnover labels to a matching model with flexible wages and efficient turnover. Quit-layoff labels are applied based on who initiates the separation by demanding a wage revision.

The model is consistent with many empirical regularities of the quit-layoff distinction. Test results based on estimates of the structural turnover model are mixed. Perhaps the most surprising feature is the very weak, negative effect of the pre-separation wage on the fraction of separations called quits.

In the face of some unsupportive evidence, one is tempted to yield to some other model or approach to account for the ill-understood behavior. That would be premature. I have shown that a simple story goes a long way. A fruitful line of research on the quit-layoff distinction would be to append the model of quit-layoff labels to more sophisticated joint-wealth-maximizing models of matching (e.g., Mortensen 1986).

I opened this chapter with a question: Why do participants in the labor market apply quit or layoff labels to employment transitions? This question is challenging if, as under the joint wealth maximizing hypothesis, there is no meaningful economic distinction between quits and layoffs. I offer one answer: the distinction or terminology is useful in summarizing the environment of the employment transition. As such, it survives in usage.

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TABLE

VARIABLE DESCRIPTION AND METHOD OF IMPUTATION

Variable	Description	No. of Missing Observations[a] 0	1	Method of Imputation
S	= 1 if changes employer	–	0	
Q	= 1 if S=1 and reason is "quit"	–	0	
L	= 1 if S=1 and reason is "laid off; fired," or "plant closing"	–	0	
E	= 1 if employed or on temporary layoff	–	0	
U	= 1 if unemployed	–	0	
WAGES	real hourly pay on main job (1972 $)	0	426	not imputed
EDUCAT	highest grade completed (to grade 17)	48	49	EDUCAT(1) = EDUCAT(0), or subsample means
ED17	= 1 if EDUCAT = 17	0	0	
EXPER	years employed since age 18[b]	80	87	$= -16.69 + 0.95\ AGE - 0.18\ EDUCAT$ (0.15) (0.002) (0.01)
TENURE	years employed with current employer	5065	7144	(see text)
GOVT	= 1 if employed by government	34	271	sample means

UNION	= 1 if union member or covered by union contract	0	0	
Industry	twelve 1-digit industry dummy variables	109	190	year by employment class means
Occupation	nine 1-digit occupation dummy variables	22	67	year by employment class means
AGE	years of age	0	0	
RACE	= 1 if white	0	0	
MARRIED	= 1 if married, spouse present	0	0	
SOUTH	= 1 if resides in South	2	4	sample means

[a] "0" denotes the initial period and "1" the subsequent period.

[b] This variable is as coded on the PSID tape; I did not increment the variable from year to year.

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