An Analysis of the Quit Rate in American Manufacturing Industry:

Preliminary Results

John Pencavel, 1968, 344 pages

The literature on voluntary labor mobility has centered around the question of the extent to which labor markets operate in accordance with the classical model. Some writers maintain that the model is seriously misleading when it is used to account for short-run market phenomena and insofar as it presumes the existence of a more or less "structureless market." On the other hand, others feel confident that the competitive model does broad justice as a basis for description of labor market behavior. As far as the study of quit rate is concerned, these investigations into the workings of labor markets have raised a number of unresolved problems which are the subject of analysis here.

1 For example, Clark Kerr writes, "Barriers to movement are set up by the skill gaps between occupations and the distance gaps between locations. Beyond the specificity of skills and the money costs of physical transfer lie such various but no less important impediments to competition as lack of knowledge, the job tastes of workers, their inertia and their desire for security, and the personal predilections of employers. The competitive market areas within which somewhat similar employers try to fill somewhat similar jobs, are normally quite restricted," The Balkanization of Labor Markets in E. Wight Bakke and others, Labor Mobility and Economic Opportunity, Wiley and M.I.T. Press, 1954, pp. 94-95.


3 Quits are terminations of employment initiated by employees for any reason except retirement, transfer to another establishment of the same firm, or service in the Armed Forces. U.S. Department of Labor, Bureau of Labor Statistics, Measurement of Labor Turnover, revised June 1965, p. 2. The quit rate is the number of quits per month for each one hundred employees.
It has been discovered that, even after taking account of the cyclical
nature of voluntary turnover, there exists a downward trend in the quit rate over
the postwar period. The attempt to explain this negative trend calls into
consideration the determinants of the quit rate. In particular, the following
questions require resolution:

1. Do individuals respond to earnings differentials by tending to leave
low wage industries or are there "no surface indications of a strong
relation between quit rate and wage level"?;

2. does the presence of trade unions significantly affect the quit rate
in an industry?;

A regression of the annual quit rate for manufacturing industries ($Q_t$) on the
civilian unemployment rate ($UN_t$) and on a simple time trend ($T_t$) spanning the
years 1947 to 1966 yields the following equation:

\[
Q_t = 4.764 - 0.443UN_t - 0.057T_t
\]

\[
(0.434) (0.094) (0.018)
\]

where the estimated standard errors are shown in parentheses below the estimated
coefficients. Both independent variables are significant and 74 per cent of the
variance of $Q_t$ is accounted for. Other equations testing for the existence of
a negative time trend are found in John E. Parker and John F. Burton, 'Voluntary
Labor Mobility in the U.S. Manufacturing Sector,' Proceedings of the Twentieth
Annual Winter Meeting of the Industrial Relations Research Association, Washington,
D.C., December 1957, pp. 61-70.

218. The question may appear beyond doubt, but one recent investigation found
in a cross-section analysis of the quit rate for the year 1963 that earnings did
not significantly (using conventional standards of significance) affect quits,
a result that was accounted for by the relatively low level of business activity
for that year. See V. Staisov and R. L. Raimon, 'Determinants of Differences in
the Quit Rate among Industries,' The American Economic Review, forthcoming.

6. The question is an important one since a recent cross-section study used the
quit rate as a proxy for the extent of unionism in an industry to examine the
degree to which unions affect relative wages over different stages of the business
cycle. See Robert V. Ealy, 'Market Power as an Intervening Mechanism in Phillips
Curve Analysis,' Economica, N.S., XXXII, 125, February 1965, pp. 48-64.
3. has the spread of benefits like pension plans and seniority rights impeded voluntary mobility?\footnote{7} and

4. is there evidence to support the proposition that "the characteristics of voluntary quitting differ markedly between those industries that are heavy users of skilled workers as contrasted with large users of lesser-skilled help"?\footnote{8}


\footnote{8} The hypothesis is advanced by Sara Behman, 'Wage Determination Process in U.S. Manufacturing,' \textit{The Quarterly Journal of Economics}, LXXXII, 1, February 1968.
These questions will be examined in this study and an explanation is offered for the postwar decline in the quit rate. Section I develops a theoretical framework for the basic empirical results in Section II. Finally Section III presents some further results and implications.

I. A Conceptual Framework

The empirical work that follows this section uses industry data. Hence the theoretical development here runs in terms of the individual choosing employment between industries rather than across occupations or geographical regions.

1. Treating the Classical theory of labor mobility in a human capital framework, it is postulated that, ceteris paribus, workers allocate their employments between industries so as to maximize their net rates of return over costs over their lifetimes. In a state of perfect information where all employment opportunities are known to him and if his employability is not restricted by lack of skills, "economic man" will move out of one employment into another if the latter offers him higher discounted real net returns. Private real returns consist of the sum of monetary earnings (taking into account any involuntary loss of employment) over the length of the working lifetime plus psychic benefits from employment in a particular occupation (or the psychic benefits from unemployment when the individual quits to leave the labor market altogether).


10. Earnings influence the decision to quit when the alternative facing the individual is not simply employment in another industry, but instead that of leaving current employment to engage in search or even quitting the labor force altogether. A significant if not the primary element in the opportunity cost of opting for either of these latter two alternatives is the level of earnings he would obtain by remaining in his current employment.
The private costs associated with quitting include the expenses involved in moving to another occupation, the income foregone while seeking out and learning a new job, and the psychic costs consequent upon leaving a familiar environment. More formally, an individual behaves according to the following expression:

$$\max_i \left[ \int_0^T Y_i e^{-rt} dt - H_i \right]$$

where $i$ denotes the industries available for employment, $Y_i$ monetary earnings in industry $i$, or the individual's subjective discount rate (assumed equal for employment in each industry), $T$ his earnings span, and $H_i$ the costs of movement into industry $i$ (which, of course, is zero for a worker currently employed in industry $i$).

2. This behavior is appropriate for the individual who, prior to quitting his current employment, has secured alternative employment in another industry. But often an employee has not ensured his future employment elsewhere before terminating his present job\(^{11}\) nor can he always be sure that business conditions will be such that he will have to select between many job offers: he feels dissatisfied with his present job and calculates that his net returns would be higher in an alternative occupation, yet he is not certain that, if he quits, he will be employed in these more lucrative jobs. In view of this employment uncertainty, it is appropriate to modify expression (2) to read:

$$\max_i \left[ p \int_0^T Y_i e^{-rt} dt - H_i \right]$$

where $p$ measures the probability that the individual will actually obtain employment in $i$. Like any other probability, $p$ lies between unity, when he has already

\(^{11}\) Lloyd Reynolds, op.cit., pp. 214-215, reported in his study of the New Haven labor market that almost 60 per cent of all voluntary quits had not lined up a new job before quitting the old one. And Charles H. Myers comments, "Workers who leave one job voluntarily for another do not usually have another job offer in mind. They may leave in the expectation that they can find a better job, but their knowledge of available alternatives is apt to be sketchy and their search haphazard," 'Labor Market Theory and Empirical Research', p. 320 of *The Theory of Wage Determination*, Macmillan, London, 1957, ed. by J.T. Dunlop.
ensured his employment before quitting his present job or when he chooses to remain in his current employment, and zero when there are no employment opportunities open to the individual owing to general slack activity or to his inadequate skills and ability. Clearly, save for the special case of \( p = 1 \), the effect of introducing employment uncertainty into the decision to quit is to reduce the attraction of alternative employments and thus reduce voluntary labor turnover. 

3. Except for the risk-neutral worker, employment preferences between industries are determined not simply by the present capital value of expected returns but also by other properties of the probability distribution of returns from available employments. In choosing between employment in two industries offering equal net returns, a risk-averting individual will prefer that industry in which these returns are fairly certain of being earned with small chance of large deviations in either direction than one in which the probability distribution of present capital values displays much greater variation. Conversely, faced with the same choice, risk-lovers are attracted by the prospect of large gains to that industry with a larger dispersion of expected returns, notwithstanding the increased chances of below-average returns that would accompany such a distribution.\(^\text{12}\) Hence the individual's utility function \( U \) includes a measure of the dispersion of returns \( \sigma_y \) around the expected value: \( U = g(\sigma_y) \). For risk-aversers, \( g' < 0 \); for risk-lovers, \( g' > 0 \); and for risk-neutral employees, \( g' = 0 \).

And the maximizing expression becomes

\[
\max_p \left[ \int_t^T e^{-rt} \ dt - H_t \ ; g(\sigma_y) \right]
\]

\(^\text{12}\) The argument is directly analogous to recent contributions in monetary economics in the theory of portfolio selection. See, for example, James Tobin, 'Liquidity Preference as Behavior towards Risk,' Review of Economic Studies, XXV, 57, February 1958, pp. 65-85.
The preceding discussion has isolated the variables pertinent to the individual's decision to voluntarily terminate his employment in one industry. The next step is to identify these determinants with observable variables. The assumption is made that each individual forms his expectation of discounted lifetime earnings in a given industry on the basis of the mean of current wages and salaries ($W_1$) and in the absence of information about the nature of the probability distribution of returns from various industries, the standard deviation ($\sigma_1$) seems the appropriate measure of dispersion. Over time, $p$, the probability that a worker who quits will find alternative employment, varies with the business cycle, but at a point in time, this probability can be made a function of the characteristics of the individual and of the nature of the particular labor market. These are considered more fully below. Finally, in moving from the determinants of an individual's decision to quit (a binary operation) to the determinants of the quit rate of an industry's labor force, it is assumed, ceteris paribus, that industry $i$'s quit rate ($Q_1$) is a linear function of the difference between earnings in $i$ and earnings in neighboring industries:

$$Q_1 = \alpha (W_1 - \bar{W})$$

where $\alpha < 0$.

The discussion so far specifies the following quit function:

$$Q_1 = \phi (W_1, \sigma W_1, P_1, L_1)$$

where $P_1$ and $L_1$ are vectors of the personal characteristics of employees and of the nature of the labor market, respectively, in industry $i$. The following paragraphs consider the specification of the vectors $P_1$ and $L_1$.

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13. The use of both the mean and the standard deviation of returns can be justified if quadratic utility functions are postulated: $U = aR + bR^2$. Taking expectations and noting that $E(R^2) - E(R)^2 = \sigma^2$ yields $E(U) = aE(R) + b[E(R)^2 + \sigma^2]$ so that $E(U) = f[E(R), \sigma^2]$.
1. The human capital treatment of mobility suggests that the longer an individual's life-span of economic activity, the higher the rate of return on an investment in changing one's industry of employment. This implies that a young labor force should display a higher quit rate than an older one.  

2. Since women spend less of their time in the labor force than do men, they tend to invest less in education and skills and this smaller investment is reflected in their lower earnings. These lower earnings imply a lower opportunity cost of their quitting employment and also of their quitting the labor force altogether. This reasoning suggests that an industry’s quit rate should be positively correlated with the proportion of women in its labor force.

3. There is considerable evidence to the effect that the propensity to quit of recently employed workers is significantly higher than for employees with longer service. The reason preferred for this relationship conventionally runs in terms of the non-wage benefits (e.g. seniority rights) that are built up with increasing attachment to a given employer, but a more general explanation can


15. V. Stoikov and R. L. Raimon, op. cit., argue that, insofar as 'females suffer discrimination in the market for labor,' their quit rate should be lower as a consequence. Clearly, discrimination in employment has the effect of reducing p and so of discouraging quits. On the other hand, when discrimination manifests itself in earnings lower than those earned by other workers, the effect is to reduce the opportunity costs of unemployment and so encourage females to quit.

16. For instance, Summer Shichter wrote, 'The concentration of the turnover among new men is not confined to any particular class of workmen. Regardless of the grade of skill of the workmen, it is the new men who most frequently leave,' The Turnover of Factory Labor, D. Appleton, New York, 1919, p. 57. Also see Reynolds, op. cit., p. 80

17. Thus, see the discussion in Farnes, op. cit., pp. 107-103.
be made by appeal to the human capital treatment of labor. Becker has argued that "Employees with specific training have less incentive to quit ... than employees with no training or general training which implies that quit rates are inversely related to the amount of specific training."\textsuperscript{18} Few jobs, if any, involve no specific training whatsoever, if only the training consists of acquainting new workers with the organization. Moreover, it is not unrealistic to suppose that the volume of specific training in an individual increases along with the duration of his employment.\textsuperscript{19} Hence the implication is that quits become less frequent as workers remain in the same employment.

4. It has been repeatedly documented that workers typically show greatest attachment to their geographical location rather than to an industry or occupation.\textsuperscript{20} Labor generally search for employment within their own community before considering moving to another region. If this is true, then the probability that an employee who quits will find alternative employment depends upon the extent of the labor market considered as a geographical unit and hence the prediction is derived that the industry's quit rate is a positive function of the proportion of its employment located in large standard metropolitan areas.

5. To the extent that job dissatisfaction encourages voluntary mobility and if "the removal of the sources of job dissatisfaction is obviously one of the prime functions, if not the prime function, of American unionism....,"\textsuperscript{21} the effect of trade unionism is to diminish quits. Movement is further discouraged


\textsuperscript{19} For instance, Paul Bissenden and Emil Frankel note that "....as the length of service of the employee increases, his value to the organization is also enhanced," \textit{Labor Turnover in Industry}, Macmillan, New York, 1922, p. 115.

\textsuperscript{20} See, for example, the discussion in Parnes, op. cit., Ch. 3.

\textsuperscript{21} Joseph Shister, \textit{op. cit.}, p. 44.
where openings are limited to union members. Notwithstanding some arguments suggesting the converse, the quit rate should be a negative function of the percentage of an industry's employment covered by collective bargaining contracts.

This section has advanced a number of hypotheses concerning the determinants of the quit rate across industries. The following sections concern themselves with the testing of these hypotheses.

II Preliminary Empirical Results

The arguments put forward in the previous section can be summarized in the following estimating equation:

$$Q_i = \alpha_0 + \alpha_1 U_i + \alpha_2 \sigma_w_i + \alpha_3 A_i + \alpha_4 F_i + \alpha_5 C_i + \alpha_6 S_i + \alpha_7 U_i + \epsilon_i$$  \( i \)

\( \alpha_1, \alpha_7 < 0; \alpha_3, \alpha_4, \alpha_5, \alpha_6 > 0; \alpha_2 \neq 0 \)  \( 23 \)

where \( Q_i \) = the quit rate in industry \( i \)

\( U_i \) = median wage and salary income (in thousands of dollars)

\( \sigma_w_i \) = standard deviation of wages and salaries (in thousands of dollars) by individuals

\( A_i \) = the proportion of employees under 30 years of age

\( F_i \) = the proportion of female employees in the work force

\( C_i \) = the accession rate lagged one year, i.e., permanent and temporary additions to employment per 100 employees  \( 24 \)

22. Parnes, op. cit., pp. 125-130, provides a readable summary of these propositions.

23. The hypothesis that \( \alpha_2 \neq 0 \) means simply that individuals may well be risk-aversers or risk-lovers, but that they are not risk-neutral.

24. This variable is meant to capture the effects of short service on the propensity to quit following the discussion in the preceding section.
SM<sub>i</sub> = the proportion of employment in large SMSAs
U<sub>i</sub> = the proportion of employment covered by collective bargaining agreements.
ε<sub>i</sub> = a stochastic disturbance term

Equation (3) was fitted to data for 40 manufacturing industries for the year 1950 using single-equation ordinary least squares (OLS) techniques and the results are shown in equation (IA) of Table I. The overall impression is that it provides a good explanation of the behavior of voluntary labor turnover. The only unambiguously insignificant coefficient is that relating to ε<sub>i</sub>, the standard deviation of wages. Not only is it insignificant, but the size of the coefficient is extremely small: an increase in the standard deviation by one thousand dollars (that is, by one-third of the mean σ<sub>ω</sub> for the whole sample of manufacturing industries) reduces the quit rate by only 0.035, a negligible amount. The hypothesis that individuals are not risk-neutral in their choice of employment between industries is not substantiated.  

25 Practically identical results are obtained by using the coefficient of variation rather than the standard deviation as the measure of wage dispersion: again the estimated coefficient is one-half its estimated standard error.

26 Considerable attention was devoted to the possibility that σ<sub>ω</sub> should be treated as simultaneously determined along with the quit rate. The proposition would run that the level of the quit rate indicates the volume of search being undertaken by employees within that industry and this should have a negative effect on wage dispersion along lines developed by George Stigler, 'Information in the Labor Market', The Journal of Political Economy, Supplement, LXX, 5, October 1962, pp. 54-105. The suspicion was not borne out: σ<sub>ω</sub> remained insignificant in the quit equation and in the second equation the hypotheses concerning the nature of the relationship between quits and wage dispersion was not indicated.
Table 1: Determinants of the Quit Rate (Q)

<table>
<thead>
<tr>
<th>Equation</th>
<th>Constant</th>
<th>C_1</th>
<th>A</th>
<th>P</th>
<th>C_2</th>
<th>U</th>
<th>ST</th>
<th>PSEE</th>
<th>R^2</th>
<th>SEE</th>
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</thead>
<tbody>
<tr>
<td>(A)</td>
<td>1.640</td>
<td>-0.266</td>
<td>-0.035</td>
<td>2.551</td>
<td>0.310</td>
<td>0.129</td>
<td>1.095</td>
<td>-0.702</td>
<td>0.770</td>
<td>0.266</td>
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<tr>
<td></td>
<td>(0.469)</td>
<td>(0.074)</td>
<td>(0.062)</td>
<td>(1.570)</td>
<td>(0.135)</td>
<td>(0.021)</td>
<td>(0.426)</td>
<td>(0.317)</td>
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<tr>
<td>(B)</td>
<td>1.576</td>
<td>-0.266</td>
<td>-0.035</td>
<td>2.314</td>
<td>0.305</td>
<td>0.126</td>
<td>0.760</td>
<td>-0.446</td>
<td>0.708</td>
<td>0.267</td>
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<tr>
<td></td>
<td>(0.461)</td>
<td>(0.073)</td>
<td>(0.063)</td>
<td>(1.562)</td>
<td>(0.134)</td>
<td>(0.023)</td>
<td>(0.421)</td>
<td>(0.336)</td>
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<tr>
<td>(C)</td>
<td>1.462</td>
<td>-0.301</td>
<td>-0.045</td>
<td>3.710</td>
<td>0.308</td>
<td>0.114</td>
<td>1.071</td>
<td>-1.006</td>
<td>0.777</td>
<td>0.270</td>
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<td></td>
<td>(0.485)</td>
<td>(0.059)</td>
<td>(0.059)</td>
<td>(1.508)</td>
<td>(0.136)</td>
<td>(0.027)</td>
<td>(0.420)</td>
<td>(0.607)</td>
<td></td>
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</tr>
</tbody>
</table>

Notes: Estimated standard errors of the estimated regression coefficients are in parentheses. SEE is the standard error of estimate for the equation.
The on the other hand, the depressing effect of the level of wages on quits seems established: an increase of one hundred dollars in annual wage income will reduce the number of quits by about 2.7 in every 100 of employees. The only variable more significant than \( W \) is \( C \), the lagged accessions rate. According to the interpretation offered above, longer service workers are rewarded for the firm's (specific) investment in them by paying them higher rewards than they could receive elsewhere hence the implication is that quit rates decline with length of service. Its healthy significance and size -- the estimated coefficient is approximately one-half of that on wages and salaries -- supports this hypothesis.

The variables \( F \), the proportion of female employees, and \( SM \), the proportion of employment in large SMSAs, are of almost equal significance, though the latter has an estimated coefficient more than three times the size of that on \( F \). Even more important in terms of the coefficient size is the variable \( A \), the proportion of employers under 30 years of age: an increase in \( A \) by a small one per cent will increase quits by almost 3 employees in every 100. Yet the statistical significance of the estimated coefficient on \( A, \hat{a}_3 \), is in question. Using the \( t \) distribution, \( \hat{a}_3 \) is not significant on a two-tail test at the 5% level, but is significant at the same significance level if one is heroic enough to apply a one-tail test.

The unionism variable, \( U \), is also significant, substantiating Slichter's belief that the lack of a grievance system "increases the importance of the grievances to the workmen and creates new grievances" so encouraging quits.\(^{27}\) It is of interest to consider whether it is not only the presence of unionism that discourages voluntary mobility, but in addition the extent to which the union is seen to be pushing for the interests of its members. The index of union activism proposed is \( ST \), the number of work stoppages averaged over the fouryears 1956 to 1959 as a percentage of total employment (in thousands).

\(^{27}\) S. H. Slichter, op. cit., p. 204.
The consequences of entering ST into regression equation (3) is shown in equation (IB) of Table I. Though it has the expected sign, it does not appear to be significant by conventional standards. Further, the inclusion of ST has the effect of reducing the size of the coefficient or unionism without simultaneously cutting its standard error. One explanation for this might be collinearity: U and ST have a simple correlation of .55. Consequently, an alternative route was found for entering ST into the quit equation. On the belief that the appropriate measure of extent of industrial jurisprudence is some composite of unionism and strike activity, a new variable was defined U*ST, the extent of unionism multiplied by the variable ST defined above. Replacing U and ST with this new variable yields the result given in equation (IC) of Table I. Comparison of equations (IA), (IB), and (IC) reveals that the coefficient estimates on all the non-union variables remain relatively stable under these different specifications. The only possible exception is the estimate for W which rises from 0.27 to 0.30 though there is no significant difference between the two estimates. The estimate for U*ST is relatively large in size and statistically significant by the conventional criteria. Since, on the basis of these estimates, it is difficult to choose between equations (IA) and (IC), the hypothesis that union militancy needs to be considered as well as the mere presence of unions is not disproved.

This section has proposed and defended a particular specification of an equation accounting for the behavior of the quit rate across industries. Of the hypotheses put forward in Section I, only that which suggests that individuals are not risk-neutral has been found lacking empirical support. The next section examines some possible biases in the above estimates and extends the analysis by considering some further hypotheses.
III Further Empirical Results

A. Questions of Simultaneity

The previous section used single-equation ordinary least squares (OLS) techniques to test the hypotheses proposed and to measure the effects of the independent variables on the dependent variable. The appropriateness of such single-equation methods might be questioned and arguments in favor of a "systems approach" advanced. In particular, developments in the theory of information and dissemination of knowledge might imply that the level of wages and salaries should be considered jointly determined with the quit rate. The argument is as follows: the employer has the choice between a low-wage, high-search strategy or a high-wage, low-search strategy; a high quit rate indicates poor matching of the employee with his job and reflects a low-search strategy on the part of the employer; hence high quit rates suggest poor selection procedures and consequently the employer must pay high wages to attract the type of labor he wants. The hypothesis may be summarized in the form:

$$W = W(Q, X)$$

where $$W Q > 0$$ and where $$X$$ is a vector of exogenous variables.

The following are the proposed arguments of $$X$$:

1. $$U_1$$ = the proportion of employment covered by collective bargaining agreements

A considerable volume of evidence has been accumulated to the effect that unions influence relative wages.28


2. \( H_1 \) = the proportion of employees who completed high school
\n\( M_1 \) = the proportion of laborers and operatives in the labor force

These variables are offered as proxies for the quality of labor and are expected to have a positive effect upon \( U \).

3. \( F_1 \) = the proportion of female employees in the work force. Insofar as wage discrimination is practiced against females and insofar as women are less skilled than men, \( F \) exerts a negative influence on \( W \).

Combining these hypotheses with equation (4) yields the following estimating equation:

\[
U_1 = \beta_0 + \beta_1 Q_1 + \beta_2 U_1 + \beta_3 H_1 + \beta_4 M_1 + \beta_5 F_1 + v_1
\]

\( \beta_1, \beta_2, \beta_3, > 0; \beta_4, \beta_5 < 0 \)

where \( v_1 \) is a stochastic disturbance term. If this line of argument is correct, estimating equation (3) by classical least-squares techniques will lead to an upward bias in the estimate of \( \alpha_1 \). So that simultaneous equation techniques are called for. Both two-stage least squares (2SLS) and three-stage least squares (3SLS) methods were used. Since the 2SLS and 3SLS estimates differ only marginally from one another, Table II presents only the more efficient 3SLS estimates. Equations (IIA) and (IIB) are the estimates of the system consisting of equations (3) and (5). Comparison of the OLS and 3SLS estimates for the quit equation reveals no significant change in the coefficients. Though \( \tilde{\alpha}_1 \) is smaller when

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30 It is not difficult to show that the asymptotic bias of the OLS estimator for \( \alpha_1 \) is

\[
\text{plim} (\tilde{\alpha}_1 - \alpha_1) = \frac{1}{\Delta} \text{var} Y \left[ \beta_1 \text{var} (\varepsilon) + \text{cov} (\varepsilon, v) \right]
\]

where \( Y \) is the vector of non-wage variables determining \( Q \) in equation (3), i.e., \( Y = (\sigma, A, F, C, S, H, U) \), and \( \Delta \) is the determinant of a positive definite matrix. Since by assumption \( \alpha_1 \leq 0 \) and \( \beta_1 > 0 \) and since we lack a priori information about \( \text{cov} (\varepsilon, v) \), \( \text{plim} (\tilde{\alpha}_1 - \alpha_1) > 0 \) save for the case in which \( R_1 = \text{cov} (\varepsilon, v) = 0 \). This argument is directly analogous to that used in O.C. Ashenfelter and G. E. Johnson, 'Unionism, Relative Wages, and Labor Quality', Osaka Economic Papers, forthcoming.
Table II: Three Stage Least Squares Estimates of the Quit (C) Equation and the Wage (W) Equation.

mean $C = 1.356$
standard deviation of $C = 0.528$

mean $W = 4.663$
standard deviation of $W = 0.850$

estimated coefficients on:

<table>
<thead>
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<th>Equation</th>
<th>Dependent Variable</th>
<th>Constant</th>
<th>$C$</th>
<th>$W$</th>
<th>$CW$</th>
<th>$A$</th>
<th>$P$</th>
<th>$C$</th>
<th>$SM$</th>
<th>$U$</th>
<th>$H$</th>
<th>$Ml$</th>
<th>$R^2$</th>
<th>$SE_R$</th>
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<tr>
<td>(IIA) C</td>
<td>1.414</td>
<td>$-0.206$</td>
<td>(0.097)</td>
<td>$-0.035$</td>
<td>3.003</td>
<td>0.363</td>
<td>0.140</td>
<td>0.797</td>
<td>$-0.830$</td>
<td>0.443</td>
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<td>(0.553)</td>
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<tr>
<td>(IIB) W</td>
<td>4.826</td>
<td>$-0.443$</td>
<td>(0.682)</td>
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<td>(0.164)</td>
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<td></td>
<td></td>
</tr>
<tr>
<td>(IIC) C</td>
<td>1.488</td>
<td>$-0.213$</td>
<td>(0.551)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>(0.097)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>(IID) W</td>
<td>3.720</td>
<td>$-0.700$</td>
<td>(0.652)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>(0.105)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

see Notes to Table I
estimated by the simultaneous equation methods, as suggested by footnote 30, the difference between the OLS and 3SLS estimate is not significant. In the wage equation (III), the hypotheses as to the effect of $U$, $H$, $M$, and $F$ are substantiated, but contrary to the previous discussion quits have a negative, not a positive affect on wages. The explanation offered for this is that the strength of the causation from $W$ to $Q$ yields such a high correlation ($-0.714$) that it even dominates in an equation in which $Q$ is the dependent and $W$ the independent variable. 32

Correlation is not causation so $Q$ is dropped from the wage equation and the system reestimated as shown in equations (IIC) and (IID). 33 Again the estimated parameters of the quit equation are to all intents and purposes unchanged. These results indicate that the conclusions arrived at in the previous section on the basis of OLS regression methods do not require amendment if quits are considered as part of a simultaneous equation system.

B. The "Industrial Feudalism" Hypothesis

The strong negative time trend in the quit rate since the war 34 has induced some commentators to speculate that the growth of pension plans and seniority rights has been an increasingly immobilizing factor in voluntary labor turnover -- in short, there is a new "industrial feudalism." 35 This hypothesis was considered

31 It is worth making it clear that the relationship here concerns the effect of the quit rate on the level of wages and does not invoke the same arguments as Sara Behman's which examine the effect of the quit rate on the rate of change of wages over time, "Labor Mobility, Increasing Labor Demand, and Money Wage-Rate Increases in United States Manufacturing", The Review of Economic Studies, XXXI (4), No. 38, October 1964, pp. 253-265.

32 In passing, those concerned with the relative wage effects of unions may be interested to note that equation (IIB) estimates that unionism has raised the average wage of union workers by some 20 per cent above that of non-union workers.

33 Equation (6) in footnote 30 shows that the use of simultaneous equation techniques is still obligatory, even if $\beta_1 = 0$, as long as $\text{cov}(\epsilon_1, \epsilon_2) \neq 0$.

34 See equation (1) in footnote 4.

35 The argument was examined and rejected by Arthur Ross, "Do We Have a New Industrial Feudalism?", The American Economic Review, XLVIII, 5, December 1958, pp. 503-520, and then reconsidered with more sympathy by John Parker and John Burton, op. cit.
by adding to equation (3) a variable measuring such supplements to wages and salaries. Though the data leave much to be desired, a heroic attempt was made to test the industrial feudalism hypothesis by forming the following three variables:

1. $SUP1 = \text{wage supplements as a proportion of wages and salaries for 1957 computed by the Department of Commerce}$
2. $SUP2 = \text{wage supplements as a proportion of wages and salaries for 1957 from the 1960 Census of Manufactures}$
3. $SUP3 = \text{employer contributions to private welfare plans as a proportion of gross payroll for 1959 from the Department of Labor}$

As is evident from Table III, the results give no support whatsoever to the industrial feudalism hypothesis. In no case are the estimated coefficients significantly different from zero and their inclusion in the regression leaves the basic structure unaltered.

If these, albeit crude, results are basically correct, then supplements to wages or fringe benefits are unlikely to have accounted for the negative trend in the quit rate. The question arises whether equation (3)'s structure can

---

35. A mere glance reveals substantial differences between the three series used, differences for which an explanation is not immediately forthcoming. Moreover, observations are available only by two-digit industry classification so a grouping of observations became necessary. This means that these statistics are for the most part only approximations to the "true" values and such measurement errors will bias (even in the probability limit) the estimated coefficients toward zero. These important limitations should not be overlooked when interpreting the results.

37. This skepticism concerning the effect of wage supplements is not particularly iconoclastic. For instance, one recent commentator has written, "...there is little evidence to support the claim that pension plans are unduly restrictive in terms of reducing labor mobility or hindering the hiring of older workers. Rather it would seem that pension plans have become the scapegoat for other more significant factors. Also, on the surface they offer an easy explanation to problems whose solution is much more complicated." Burt K. Scanlon, 'Effects of Pension Plans on Labor Mobility and Hiring Older Workers', Personnel Journal, 44, 1, January 1965, p. 34.
### Table III: The "Industrial Feudalism" Hypothesis

<table>
<thead>
<tr>
<th>Equation</th>
<th>Constant</th>
<th>$W$</th>
<th>$\sigma_w$</th>
<th>A</th>
<th>F</th>
<th>C</th>
<th>SM</th>
<th>U</th>
<th>SUP1</th>
<th>SUP2</th>
<th>SUP3</th>
<th>$R^2$</th>
<th>SEE</th>
</tr>
</thead>
<tbody>
<tr>
<td>(IIIA)</td>
<td>1.554</td>
<td>-0.255</td>
<td>-0.034</td>
<td>2.913</td>
<td>0.305</td>
<td>0.123</td>
<td>1.014</td>
<td>-0.650</td>
<td>-0.132</td>
<td></td>
<td></td>
<td>0.778</td>
<td>0.273</td>
</tr>
<tr>
<td></td>
<td>(0.511)</td>
<td>(0.075)</td>
<td>(0.070)</td>
<td>(1.655)</td>
<td>(0.141)</td>
<td>(0.028)</td>
<td>(0.431)</td>
<td>(0.334)</td>
<td>(1.004)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>(IIIB)</td>
<td>1.814</td>
<td>-0.246</td>
<td>-0.020</td>
<td>2.524</td>
<td>0.272</td>
<td>0.127</td>
<td>1.053</td>
<td>-0.585</td>
<td>-3.246</td>
<td></td>
<td></td>
<td>0.782</td>
<td>0.270</td>
</tr>
<tr>
<td></td>
<td>(0.534)</td>
<td>(0.077)</td>
<td>(0.071)</td>
<td>(1.664)</td>
<td>(0.143)</td>
<td>(0.020)</td>
<td>(0.428)</td>
<td>(0.345)</td>
<td>(3.740)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>(IIIC)</td>
<td>1.525</td>
<td>-0.270</td>
<td>-0.038</td>
<td>3.054</td>
<td>0.316</td>
<td>0.125</td>
<td>1.006</td>
<td>-0.700</td>
<td></td>
<td>0.272</td>
<td></td>
<td>0.778</td>
<td>0.273</td>
</tr>
<tr>
<td></td>
<td>(0.516)</td>
<td>(0.082)</td>
<td>(0.074)</td>
<td>(1.347)</td>
<td>(0.147)</td>
<td>(0.025)</td>
<td>(0.430)</td>
<td>(0.322)</td>
<td>(2.460)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

see Notes to Table I.
explain this trend without resort to variables like wage supplements. The basic procedure is to enter into the estimated 1959 equation values for the variables for all manufacturing industry for given years to see whether this structure can account for the trend behavior in the quit rate.

First, since there are available data for only five non-census years, an estimate of the trend over each of these years is required. Hence the following regression equation was estimated from annual data spanning the period 1947 to 1958:

\[
Q_t = 4.076 - 0.464 \text{ UN}_t - 0.044 \text{ T}_t - 0.050 \text{ D}_{53-56} - 0.051 \text{ D}_{55-56} \\
(0.531) (0.113) (0.010) (0.038) (0.036) (0.030)
-0.031 \text{ D}_{58-59} - 0.046 \text{ D}_{50-52} \\
(0.025) (0.024)
\]

(7)

where \(Q_t\) is the manufacturing quit rate and \(\text{UN}_t\) is the civilian unemployment rate. \(\text{T}_t\) is a time trend taking sequential values outside the 1953-1952 period. \(\text{D}_{53-56}\) is the time trend for the few years from 1953 to 1955, \(\text{D}_{55-56}\) is the time trend for the three years 1955 to 1958, \(\text{D}_{50-50}\) is the time trend for the three years 1950 to 1952, and \(\text{D}_{50-52}\) is the time trend over the years 1950 to 1952.

Hence the estimated coefficients on the \(D\) variables measure the average decline in the quit rate over the years indicated after taking account of the cyclical nature of the quit rate. It is these estimated coefficients that our cross-section structure will try to explain.

The next step is to reestimate equation (3) for the year 1959 by omitting or replacing variables for which observations are lacking in non-census years.

The reestimated equation is as follows:

\[
Q_i = 2.270 - 0.233 \text{ UN}_i + 1.739 \text{ A}_i + 0.253 \text{ F}_i + 0.146 \text{ C}_i - 0.729 \text{ SH}_i - 0.848 \text{ U}_i \\
(0.521) (0.058) (1.445) (0.242) (0.122) (0.025) (0.295)
\]

(8)

\[R^2 = 0.752 \quad \text{SEE} = 0.250\]
where \( w_i \) has been omitted and where \( SH_i \), the proportion of employees located in the South, has replaced \( SH_i \). Comparison of this equation with equation (IA) reveals no significant change in the estimated coefficients. Entering into equation (8) changes in the values of the independent variables over the years 1955 to 1958 yields the change in the quit rate for this period as predicted by the cross-section equation. The same process is repeated to derive predictions for the quit rate change between 1955 and 1958, 1958 and 1960, and 1960 and 1962.

The final step is to compare these predicted changes in the quit rate with the trend in the quit rate over these years. The estimated coefficients on the variables \( D \) measure the trend per year so must be multiplied by the number of years covered to get the total change in the quit rate. Table IV compares the trend decline in the manufacturing quit rate during each of the four periods as derived from equation (7) with the change in quits predicted from equation (8).

The third column shows the difference between the actual trend values and the predicted values.

38. The zero-order correlation coefficient between \( SH \) and \( SN \) is = 0.501.

39. Values for \( F \), \( C \), \( SH \), and \( U \) were obtained without difficulty. Values for \( A \) were calculated from a time trend between the 1950 and 1960 Census years. Finally, \( W \) measures the average annual earnings per full-time employee in manufacturing and was deflated by a consumer price index with 1955 = 100. This implies that over time, as real incomes rise, the quit rate falls towards some limit, an implication not inconsistent with the following time series (1947-1966): regression where the trend appears as a quadratic and its coefficient has a positive sign:

\[
Q_t = 5.392 - 0.359 \text{ UH}_t - 0.317 T_t + 0.012 T_t^2
\]

\[
(0.241) \quad (0.046) \quad (0.033) \quad (0.002)
\]
Table IV: Explaining the Trend on the Quit Rate

<table>
<thead>
<tr>
<th></th>
<th>change in quit rate</th>
<th>change in quit rate</th>
<th>difference between actual and predicted values</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>predicted from</td>
<td>caused by time trend</td>
<td></td>
</tr>
<tr>
<td>equation (8)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>1953-56</td>
<td>-0.359</td>
<td>-0.361</td>
<td>0.002</td>
</tr>
<tr>
<td>1956-58</td>
<td>-0.172</td>
<td>-0.154</td>
<td>-0.018</td>
</tr>
<tr>
<td>1958-59</td>
<td>-0.034</td>
<td>-0.094</td>
<td>0.060</td>
</tr>
<tr>
<td>1960-62</td>
<td>-0.105</td>
<td>-0.176</td>
<td>0.070</td>
</tr>
</tbody>
</table>

The conformity of the predicted decline in the quit rate with the estimated trend decline is striking for each of the four periods. This impression is confirmed by following Christ's suggestion and calculating a more conservative estimate of the standard error of forecast, namely \( \hat{s} = s/\sqrt{N+1} \) where \( s \) is the unbiased estimator of the disturbance standard error and \( N \) is the size of the sample.\(^{40}\) From regression equation (8), \( \hat{s} = 0.260 \), a figure that easily covers the forecast errors by a very wide margin. This indicates that structural changes within manufacturing appear to explain the fall in the quit rate over the past 20 years. In particular, rising real earnings, the aging of the manufacturing labor force -- in 1950, 28.3 per cent of manufacturing employment was under 30 years of age, but in 1960 the corresponding figure was 10.5 per cent -- and appropriate movements in the lagged accessions rate have been the main factors.

\(^{40}\) Christ's argument can be summarized as follows. From the multiple regression equation \( Y = \alpha + \beta X + \epsilon \), the estimator of the forecast variance \( \sigma_f^2 \) is provided by

\[ \sigma^2 \left( \frac{1}{N} + \frac{1}{X^\prime (X^\prime X)^{-1} X} \right) \]

where \( X \) is the vector containing observations on the independent variables in the prediction period. Since the middle term (a quadratic form) in parentheses is nonnegative definite for all \( X \), \( \sigma^2 \geq \sigma^2 (X + 1) \approx \sigma^2 \) and substituting \( \sigma^2 \) for \( \sigma_f^2 \) provides an easily computable statistic for testing predictions, even though the probability of a type I error is larger than necessary. See Carl F. Christ, *Econometric Models and Methods*, Wiley, New York, 1956, pp. 556-557.
generating the predicted decline. Resort to explanations like the spread of pension plans and seniority rights do not seem to be necessary to account for the negative trend in the quit rate in manufacturing industry.

C. The incidence of the Quit Rate by Occupation

The propensity to quit by occupation has been the subject of considerable investigation. Early work found that turnover rates were much higher for unskilled workers, but it was not made clear whether or not this was the product of a higher incidence of involuntary turnover among the less skilled. More recent studies, on the other hand, have implied that the opposite relationship may be nearer to the truth. To examine the effect of the occupational structure of an industry's work force on its quit rate, the following occupational mix variables were added to regression equation (3):

1. M1 = the proportion of operatives and laborers in employment.
2. M2 = the proportion of managers, salesmen, professional workers, and craftsmen in employment.
3. M3 = the proportion of craftsmen in employment.

The results are shown in Table V. For none of the indices of occupational mix is the estimated standard error larger than the estimated coefficient. The similarity of equation (VA) with (VB) reflects the high correlation between M1 and M2. These negative results indicate that for this sample the occupational structure of the industry appears to have a negligible influence on voluntary turnover. This does not support Behm's statement that "the characteristics

41. See, for instance, Slichter, op. cit., pp. 57-74.
42. Lloyd Reynolds, op. cit., writes "...movement is likely to be greatest in the skilled craft occupations", p. 75.
43. The question arises whether this sample has sufficient variability in its occupational structure to provide an adequate test of the occupational mix hypotheses. The coefficients of variation for M1, M2, and M3 for this sample of 45 manufacturing industries are 0.255, 0.265, and 0.348 respectively. These values are well within the range of coefficients of variation for the other included variables -- from 0.109 for V to 1.275 for F.
Table V: The Quit Rate and the Occupational Structure of the Industry

<table>
<thead>
<tr>
<th>Equation</th>
<th>Constant</th>
<th>$w$</th>
<th>$\sigma_w$</th>
<th>$A$</th>
<th>$F$</th>
<th>$C$</th>
<th>$SM$</th>
<th>$U$</th>
<th>$M_1$</th>
<th>$M_2$</th>
<th>$M_3$</th>
<th>$R^2$</th>
</tr>
</thead>
<tbody>
<tr>
<td>(VA)</td>
<td>2.003</td>
<td>-0.310</td>
<td>-0.050</td>
<td>3.261</td>
<td>0.295</td>
<td>0.121</td>
<td>0.965</td>
<td>-0.616</td>
<td>-0.338</td>
<td></td>
<td></td>
<td>0.780</td>
</tr>
<tr>
<td></td>
<td>(0.797)</td>
<td>(0.107)</td>
<td>(0.074)</td>
<td>(1.649)</td>
<td>(0.139)</td>
<td>(0.031)</td>
<td>(0.434)</td>
<td>(0.353)</td>
<td>(0.579)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>(VB)</td>
<td>1.686</td>
<td>-0.306</td>
<td>-0.047</td>
<td>3.146</td>
<td>0.296</td>
<td>0.123</td>
<td>0.971</td>
<td>-0.625</td>
<td>0.345</td>
<td></td>
<td></td>
<td>0.780</td>
</tr>
<tr>
<td></td>
<td>(0.505)</td>
<td>(0.103)</td>
<td>(0.072)</td>
<td>(1.608)</td>
<td>(0.139)</td>
<td>(0.030)</td>
<td>(0.433)</td>
<td>(0.349)</td>
<td>(0.616)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>(VC)</td>
<td>1.644</td>
<td>-0.275</td>
<td>-0.034</td>
<td>3.020</td>
<td>0.311</td>
<td>0.125</td>
<td>1.014</td>
<td>-0.701</td>
<td>0.170</td>
<td>0.778</td>
<td>0</td>
<td>0.636</td>
</tr>
<tr>
<td></td>
<td>(0.499)</td>
<td>(0.082)</td>
<td>(0.069)</td>
<td>(1.592)</td>
<td>(0.137)</td>
<td>(0.032)</td>
<td>(0.429)</td>
<td>(0.321)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>
of voluntary quitting differ markedly between those industries that are heavy
users of skilled workers as contrasted with large users of lesser-skilled help. 44
But it does not contradict her belief that the mobility of skilled workers is low
when business conditions are slack, but is highly responsive to buoyant demand.
Since the economy was experiencing an average civilian unemployment rate of 5.5%
in 1959, the level of business activity does not fulfill Behman's condition for
skilled workers to be the highly mobile group.

IV Conclusions
Subject to the important qualification that some of the variables employed
in the preceding statistical results are plagued with measurement error, it has
been demonstrated that many of the prevailing hypotheses concerning the operation
of labor markets have substance. In particular, the Classical model's predic-
tions concerning the responsiveness of labor to wage differentials and some
predictions derived from the human capital approach to labor market behavior are
vindicated. Also the tentative finding is that the fears of a "new industrial
feudalism" are groundless, that the post-war decline in the quit rate can be
accounted for by structural changes within the manufacturing sector.

44 Sara Behman, 'Wage-Determination Process... ', op. cit., p. 121. For her
purposes, the appropriate occupational group that measures the amount of skilled
labor is the category craftsmen which is measured by our variable N3.