The Quit Propensity of Married Men
Author(s): Kathryn L. Shaw
Source: *Journal of Labor Economics*, Vol. 5, No. 4, Part 1 (Oct., 1987), pp. 533-560
Published by: The University of Chicago Press on behalf of the Society of Labor Economists and the National Opinion Research Center.
Stable URL: http://www.jstor.org/stable/2534940
Accessed: 16/01/2009 15:41

The Quit Propensity of Married Men

Kathryn L. Shaw, *Carnegie-Mellon University*

This paper hypothesizes that the quit propensity of married men rises with an increase in their wives' income. Assuming that individuals are risk averse and that quitting is risky, the wife's income increases the husband's expected value of quitting by reducing the variance of expected family income. Using the longitudinal data from the Michigan Panel Study of Income Dynamics (PSID), the wife's income is found to have a large effect on quits. The average husband's quit rate increases by about 45% when the wife's income rises from zero to two-thirds that of the husband's. The wife's income effect nearly offsets the negative effect that marriage typically has on male quit rates.

Empirical models of quit behavior have produced three stylized facts. The propensity to quit declines with employee tenure, with age, and with marriage. The first two relationships are consistent with the hypotheses of an extensive theoretical literature focusing on firm-specific

I wish to thank Robert Avery, Dennis Epple, Steven Garber, V. Joseph Hotz, Jane Mather, Guilherme Sedlacek, Kenneth Singleton, and members of the Carnegie-Mellon Empirical Workshop and the Federal Reserve Board seminar for their comments.

[*Journal of Labor Economics,* 1987, vol. 5, no. 4, pt. 1]

0734-306X/87/0504-0002$01.50

human capital investment and/or the implicit contractual relationship between the employee and employer. The theoretical importance of tenure and age, as well as their sizable quantitative effects, warrants the extensive emphasis on these variables. However, the quantitative effect of marriage is also very sizable: the average quit rate for married men aged 25–54 is 7.4%, compared to 15.4% for unmarried men.[1] To address the well-known effect of marriage, a marriage dummy variable is generally included in empirical analysis of quit behavior. However, the cause of the negative effect that marriage has on male quit rates is often described in terms of vague sociological relationships, such as the "stabilizing influence" of marriage (Viscusi 1979*b*, p. 394). Recent increases in the labor force participation rate of married women and the concomitant emphasis on the economic role of women in the family suggest that the marriage effect needs to be analyzed from an economic perspective.

This paper specifically models the effect of the wife's income on the husband's quit rate, developing and testing the hypothesis that an increase in the wife's income will increase the quit propensity of the husband. This hypothesis is formulated by treating human capital wealth much like financial wealth. The portfolio allocation model of financial wealth is loosely applied to the quit decision after redefining wealth as the sum of the human capital of the husband and the wife. Households allocate their human wealth across alternative jobs to maximize expected utility. The husband's decision to quit is then characterized as the decision to allocate his human capital wealth to an option that is riskier than remaining in his current job. Assuming risk-averse preferences, the decision to quit is more probable when the wife is working because total household wealth can be divided between the risky quit option and her less risky current job. When the wife is not working, quitting places all wealth in the riskier option. In addition, if there are high transactions costs associated with quitting, such as unemployment, or if the household is liquidity constrained, the wife's income could assist in the financing of these transactions costs. Given equal expected wage rates on alternative jobs for husbands with and without working wives, those with working wives would have a greater expected value to quitting.

The hypothesized positive effect of the wife's earnings on the husband's quit propensity is examined empirically, and the results are very consistent with the hypothesis. Using longitudinal data from the Michigan Panel Study of Income Dynamics (PSID), the coefficient on the wife's share of household income is strongly positive in the quit equation, holding constant other standard determinants of quits. The effect nearly offsets the negative effect that marriage has on the propensity to quit. For an

[1] These are average quit rates for men in the PSID subsample to be discussed later.

average married male between the ages of 30 and 49, the predicted quit rate rises from 4.4% to 6.4% when the wife's contribution to household earnings increases from 0% to 40%. Thus the predicted quit rate for a married male with a working wife approaches the predicted quit rate of an unmarried male, which is 7.1%.

Section I below develops the portfolio allocation model of the propensity to quit for married men and includes a discussion of the riskiness of quitting. This model extends the literature that emphasizes the effects of risk aversion on human capital investment. In Section II the empirical specifications of the model are laid out. The empirical results are presented in Section III, followed by the conclusions and the implications for the trend growth of the aggregate quit rate.

I. Theoretical Framework

Quits are defined to be employee-initiated job separations arising in the process of job search. An individual quits when the expected value of alternative employment surpasses the expected value of the current job. In forming these expected values, individuals are assumed to be risk averse, and quitting is assumed to be inherently "riskier" than staying on the current job. One option is riskier than another if the variance of the benefits facing the individual is greater in that option. These assumptions, as described below, shape the hypotheses of the model.

An individual may quit either with an accepted alternative job offer or without one. When an individual quits without a specific alternative job offer, the decision is clearly riskier. It is riskier because the outcome of the draws from the wage offer distribution is unknown and because there is a potential unemployment associated with search. About half of all quitters suffer some unemployment, as estimated below. In contrast, quits that occur with a specific accepted alternative wage offer would appear to be risk free. In fact, they are still likely to be relatively risky. The recent job matching literature is built on the hypothesis that alternative jobs with specific starting wage offers are still characterized by considerable uncertainty (see Johnson 1978; and Viscusi 1979*a*). The life-cycle profile of wage rates at the alternative firm depends on the quality of the "match" between the firm and the individual, which becomes apparent only after the individual begins work on the job. Thus quitting is always risky, even with an accepted alternative wage offer.[2]

As this discussion implies, quitting is distinctly different from being laid off. In theory, the two need not be distinguished. By definition, a

[2] Matching models also imply that the returns to the current job are uncertain because the quality of the match is revealed slowly over time. However, given continuous learning and updating of priors, quitting is generally riskier than staying on the current job.

layoff occurs when the worker's value of marginal product falls below his wage rate. Thus the employer could effect a layoff by lowering the worker's wage rate, thereby inducing him to quit. However, because of various institutional features of the labor market, layoffs tend to be employer-initiated terminations that occur without a decrease in the wage rate. Thus the termination decision is not the result of an optimal decision on the part of the employee. He is not given the option to choose to stay in the current job at a lower wage rate, and thus separation cannot be considered a voluntary optimizing decision.

Turning now to the application of the portfolio model, the decision process underlying the quit decision can be described in terms of the optimal allocation of human capital wealth. The household has human capital wealth equal to $K_h + K_w$, where K_h and K_w are the stocks of the human capital wealth of the husband and the wife, respectively. They have a set of alternative jobs to choose from that will earn returns on their wealth. Assume for the moment that there are no constraints on the number of jobs each individual can hold and assume that only monetary returns differ across jobs. Each individual has a vector of alternative jobs that are analogous to the "assets" of the financial allocation literature and are represented by $A_h = (A_{h1}, \ldots, A_{hJ})$ and $A_w = (A_{w1}, \ldots, A_{wJ})$, where A_{ij} is the value of human wealth placed in the jth job (for $j = 1, \ldots, J$, job options), and $i = h, w$, for the husband and wife. These jobs all earn stochastic rates of return to current wealth, $r_{h1}, \ldots, r_{hJ}$, and $r_{w1}, \ldots, r_{wJ}$, equal to the present value of each job option. These rates of return include differences in life-cycle wage growth and the future growth of human capital wealth.

The household's choice problem is to allocate wealth across assets to maximize expected utility, subject to the constraints that allocations sum to wealth and are nonnegative. However, the Kuhn-Tucker conditions resulting from the maximization of a general utility function cannot be solved analytically without making simplifying assumptions regarding the nature of the utility function and the subjective joint probability distribution of rates of return. For this reason, empirical models of financial asset choice frequently assume that investors' preferences are a function of only the mean and variance of returns.[3] The mean-variance

[3] The mean-variance approach is not very desirable because it can be shown that quadratic utility—with increasing absolute risk aversion—induces mean-variance preferences in expected utility maximization, or that rates of return are normally distributed, or that other special assumptions induce mean-variance. However, Samuelson (1970) has shown in his Fundamental Approximation Theorem that mean-variance analysis can be interpreted as an approximation to arbitrary expected utility maximization when the risk is small ($\sigma^2 \rightarrow 0$), which would be likely when holding periods are short or in a continuous time model of trading.

specification is particularly attractive in empirically implementing a model of quit behavior because alternative returns to investment enter the model in a simple quadratic manner. Given the researcher's inability to observe rates of return to alternative jobs, parameter estimation based on a highly nonlinear model that imposes a specific utility function, and therefore relies on estimated individual rates of return, is likely to be less informative. The mean-variance model is more amenable to the reduced-form estimation of rates of return.

The general mean-variance objective function is[4]

$$\max_{a} a'\bar{r} - (\rho/2)a'\boldsymbol{\Omega}a \tag{1}$$

subject to

$$\mathbf{1}'a \leq 1, \quad a \geq 0,$$

where a is the vector of asset shares, $a \equiv (A_h/K_o,\ A_w/K_o)$ for wealth $K_o \equiv K_h + K_w$; $\bar{r}$ is the vector of expected returns; $\boldsymbol{\Omega}$ is the variance-covariance matrix of returns; and ρ is the parameter of constant relative risk aversion. This particular mean-variance specification is chosen because it arises from Samuelson's approximation to expected utility maximization (see n. 3 above) and because portfolio shares are independent of initial wealth (relaxed below).

The mean-variance model will be applied to the quit decision to suggest that the allocation of human capital wealth is similar to the allocation of financial wealth, but, first, two severe restrictions are required. The first is the reduction of the asset share vector; the husband may allocate his human capital to only two options, quitting or staying on the current job, while, for simplicity, the allocation of the wife's human capital is assumed to be fixed.[5] The resulting assest share vector

[4] The use of income instead of consumption in the utility function can cause severe problems when analyzing risk aversion because consumption can be smoothed over time considerably more than income—see Hause (1974). However, I am not going to be concerned with the variation in alternative income streams over time but instead am concerned only with variation in the expected present value of returns. Thus I am assuming that utility is a function of the rest of lifetime income or that any variability in income over time could be smoothed perfectly with the capital markets. To the extent that income variation is an important part of the quit decision, particularly when human capital investment is endogenous, it will be picked up empirically by age and experience variables.

[5] The assumption that the wife's allocation is fixed is made for two reasons. First, it is highly unlikely that both the husband and the wife would quit simultaneously, except in the case of geographic migration. The amount of migration is quite small relative to quits; the results are unchanged when migrants are excluded from the empirical analysis. Second, the wife's labor force participation decision may well change when the husband is at the margin in

is $a \equiv (\alpha_q, \alpha_s, \alpha_w)$, where q denotes the allocation of the husband's human wealth to the quit option, s denotes allocation to the stay option, and α_w is the wife's initial share of the household's human capital wealth that is allocated to the market. The husband's return to the quit option equals the expected value of quitting based on an ex ante set of alternative jobs, 1, . . . , J, from which only one job is chosen ex post.

The second restriction on the model is the assumption that only the husband's quit option is subject to uncertainty, or that the variance-covariance matrix of returns degenerates to σ_q^2, the variance of the return to quitting. Certainly, the returns to staying on the current job, for the husband and wife, are also stochastic. However, it is reasonable to assume that they are considerably less uncertain or that their variance is considerably smaller than the variance of the return to quitting. More important, the econometrician has a very limited ability to model the potential covariance structure between the husband's options and the wife's options. The covariances between the two vectors of the rates of return to the husband's and wife's skills are a function of the characteristics of all their alternative jobs, which are not readily forecasted. Thus the quit model retains the mean-variance specification but reduces to the decision to allocate human capital wealth between the risky quit option and riskless option of staying, conditional on the wife's labor market allocation of human capital wealth.

Before solving this simplified model, there are some important additions, namely, the inclusion of several types of transactions costs. (1) Monetary transactions costs are generally incurred in moving from the current job to an alternative. They can be very high and uncertain if they include the cost of potential unemployment as well as direct search costs. (2) Families are frequently liquidity constrained. Liquidity constraints are defined here to be the limited ability to turn either human capital wealth or financial wealth into liquid assets that could be used to finance current transactions costs. This limited ability to finance upfront transactions costs by borrowing on future returns to human capital will increase the value of the riskless asset, or the wife's income. Her income can be used to finance transactions costs, either directly or indirectly as collateral. The effect of the wife's income is analogous to the familiar added-worker hypothesis when female labor force participation is made endogenous. (3) Families frequently incur debt, the most obvious example

making his quit decision, but to introduce the wife's labor force participation decision directly into the theoretical portfolio allocation model would require that nonmarket work be treated as an alternative option or asset. In so doing, it would be necessary to add the value of leisure to the objective function and thus very considerably complicate the theoretical model. The potential endogeneity of the wife's decision is treated in the empirical work here.

being a home mortgage. When fixed obligations are incurred, the cost of bankruptcy rises and the willingness to assume risky options may decline. Thus the cost of the higher probability of bankruptcy associated with quitting should be deducted from the returns to quitting. These additions are collectively called the stochastic transactions costs associated with the quit decision. They are incorporated into the asset choice decision by specifying costs as a function of the asset share, $C = c\alpha_q$, and thus subtracting c from the return to quitting.

The desired share of household human wealth allocated to quitting, derived from the maximization of equation (1) subject to the above additions and restrictions, reduces to

$$\alpha_q^* = (\bar{r}_q - r_s - \bar{c})/\rho(\sigma_q^2 + \sigma_c^2 + 2\sigma_{qc}), \tag{2}$$

where $\bar{r}_q$ is the expected return to quitting, r_s is the certain return to staying, $\bar{c}$ are the expected transactions costs, σ_q^2 is the variance of the return to quitting, σ_c^2 is the variance of c, and σ_{qc} is the covariance of r_q and c. The covariance would be positive if greater search increased the value of the wage offers.

The desired asset share of (2) can now be utilized to illuminate the nature of the standard qualitative quit decision. Generally, the husband must quit or not quit his primary job rather than allocate human capital wealth simultaneously across two options, the current job and an alternative quit option. The qualitative quit decision evaluates the levels of expected utility at two alternative corner solutions, quit or not quit. Using the mean-variance objective function (1) to evaluate each corner solution, the individual quits when the expected value of quitting exceeds that of staying, or

$$Q = \begin{cases} 1 & \text{if } (\bar{r}_q - r_s - \bar{c})/\rho\sigma^2 - (1/2)(1 - \alpha_w) > 0, \\ 0 & \text{if } (\bar{r}_q - r_s - \bar{c})/\rho\sigma^2 - (1/2)(1 - \alpha_w) \leq 0, \end{cases} \tag{3}$$

where $(1 - \alpha_w)$ is the husband's initial share of household wealth,

$$\sigma^2 \equiv \sigma_q^2 + \sigma_c^2 + 2\sigma_{qc},$$

and $Q = 1$ symbolizes the decision to quit. To clarify the intuition underlying (3), substitute the desired continuous asset share, α_q^*, from equation (2) into the quit decision (3) to get

$$Q = \begin{cases} 1 & \text{if } \alpha_q^* - (1/2)(1 - \alpha_w) > 0, \\ 0 & \text{if } \alpha_q^* - (1/2)(1 - \alpha_w) \leq 0. \end{cases} \tag{4}$$

Quit equation (4) implies that the probability of quitting increases when the desired share of wealth allocated to the quit option increases relative to the husband's total initial share of household human wealth, $(1 - \alpha_w)$. Or, in terms of the wife's wealth, the qualitative choice model implies that increases in the wife's wealth, α_w, will increase the husband's propensity to quit. The household could attain a higher level of expected utility if human capital wealth could be allocated continuously across two or more jobs to select the preferred level of portfolio risk. In some cases, multiple job choices can occur. For example, in considering self-employment, an individual would be likely to start his new firm in his leisure hours because the negative effect of risk aversion on expected utility is very large given the sizable degree of uncertainty prior to self-employment. However, quitting is generally considered to be solely a discrete option, in which case the wife's income increases the husband's quit propensity by decreasing the desired amount of the family's human capital wealth that is placed in the risky quit option.

Before proceeding to the estimation of equation (3), the implications of relaxing the assumption of constant relative risk aversion should be addressed. Constant relative risk aversion is pervasive in the finance literature because it is so useful. Removing this assumption results in asset demand equations that are highly nonlinear functions of initial wealth, varying across individuals and resulting in a wealth effect. A wealth effect implies that the share of wealth allocated to risky assets increases with initial wealth, given decreasing relative risk aversion, or that the level of wealth allocated to risky assets increases, given declining absolute risk aversion. Weiss (1976) used this framework to demonstrate that greater initial wealth should increase the probability of choosing a risky occupation. Analogously, greater household wealth would increase the probability of quitting in the presence of declining risk aversion, all else constant. Because the wife's income increases the absolute value of the household's monetary wealth, her income could have a wealth effect on the quit propensity. This possibility is tested in Section III.[6]

II. Empirical Specification of the Model

The empirical implementation of the quit model requires that proxies be developed for the unobservables in equation (3). Following the specification of these proxies, two methods of correcting for potential simultaneity and selectivity in the wife's share of wealth are discussed.

The quit equation (3) is entirely a function of unobserved variables,

[6] There is also a second wealth effect associated with increases in the wife's income. It has been shown (Lippmann and McCall 1976) that an increase in household wealth increases search and thus increases the expected gains to search. This possibility is included in the empirical tests for a wealth effect.

including the difference between the expected rates of return, $\bar{r}_q - r_s$, and the expected transactions costs of quitting, $\bar{c}$. These unobservables are both assumed to be a function of sets of observed variables:

$$\bar{r}_q - r_s = X'_1 B_1 + e_1, \tag{5a}$$

$$\bar{c} = X'_2 B_2 - b a_w + e_2, \tag{5b}$$

where X_1 is a reduced-form vector of personal and market characteristics that reflect the difference in the mean return to quitting versus not quitting, as described with the empirical results. Vector X_2 contains the determinants of transactions costs, which are a function of the potential duration of unemployment, the area unemployment rate, individual schooling, measures of household debt (home mortgage), and liquidity constraints. The coefficient b on the wife's share of initial household income, a_w, represents the effect that the wife's income may have in financing the costs of search.[7] In these equations, and all those below, all variables vary over time and across individuals. For ease of notation these subscripts are not shown. All parameters are constant across time and individuals, unless otherwise noted. The error terms e_1 and e_2 are identically and independently distributed (i.i.d.) $N(0, \sigma_1^2)$ and $N(0, \sigma_2^2)$ (to be relaxed below).

Inserting (5a) and (5b) into (3) and rearranging, the quit equation becomes

$$Q = \begin{cases} 1 & \text{if } X'B - \rho\sigma^2 + \rho\sigma^2 a_w + b a_w > \varepsilon, \\ 0 & \text{if } X'B - \rho\sigma^2 + \rho\sigma^2 a_w + b a_w \leq \varepsilon, \end{cases} \tag{6}$$

where $X'B = X'_1 B_1 - X'_2 B_2$ (and all coefficients B and b are multiplied by 2), $\varepsilon = e_1 - e_2$, and a_w is the wife's fixed initial share of household income, which is the proxy for the wife's initial share of household human capital wealth. Given the assumption that ε is distributed normally, the estimation technique is probit. When ε is interpreted as

[7] The wife's share of income instead of her level of income is used in the transactions costs function because the share has intuitive appeal, aside from the desire to conform to the specification used in the mean-variance model. Families are likely to incur fixed obligations that are proportional to the size of their budget, such as a home mortgage. In this case, the effect of the wife's level of income will vary with the size of the budget or with the husband's income. For example, a fixed level of the wife's income, such as $10,000, will have a much greater effect in offsetting transactions costs when the husband's income is equal to the wife's than when his income is twice that of the wife's. Thus the objective function in eq. (1) was also selected, in part, because the share formulation more closely measures the family's ability to absorb income variability.

the error in the expected net gains to quitting, the i.i.d. assumption on the error term can be replaced by the assumption that

$$\varepsilon_{it} = \phi_i + u_{it},$$

$$u_{it} = \pi u_{it-1} + \mu_{it}, \tag{7}$$

$$\mu_{it} \text{ i.i.d. } N(0, \sigma_\mu^2),$$

where i is the individual, and t is time.

In equation (6), the coefficient on the wife's share of household income is equal to the degree of constant relative risk aversion times the variance of the net return, $\rho\sigma^2$, plus the effect of liquidity constraints, b. The positive effect of $\rho\sigma^2 a_w$ on the quit propensity is very intuitive. When individuals are more risk averse, or when ρ rises, the desired share of wealth placed in the quit option falls, as described by (2). Therefore an increase in the wife's income is needed to induce a quit. The net effect is that the wife's income is more important when relative risk aversion is high, so the coefficient on her income share rises with risk aversion in the quit equation. If preferences are not risk averse, her income share would be irrelevant (disregarding transactions costs). The coefficient on the wife's share in the quit equation also includes the variance of net returns, σ^2, which has an interpretation similar to that of ρ.

Given this theoretical framework, the assumption of constancy of the coefficient on the wife's income share would be severely restrictive. The degree of constant relative risk aversion and the variance of alternative wages certainly vary across individuals, implying a model with varying slope parameters. If the specification of these is nonstochastic, or if

$$\rho = Z_1' P_1, \tag{8a}$$

$$\sigma^2 = Z_2' P_2, \tag{8b}$$

then interaction terms are introduced into equation (6), where Z_1 and Z_2 are sets of variables interacting with the wife's share of income. Since there should be considerable overlap in these sets of variables, it will not be easy to separate the determinants of ρ and σ^2.

Thus far, the wife's income share has been assumed to be exogenous to the quit decision, thereby reflecting the effect of risk aversion. Her income share may also represent an endogenous labor supply reaction to the husband's quit propensity. The wife may be induced to enter the labor force or work longer hours if either her income is needed to finance the expected transactions costs associated with his quitting (the added-worker hypothesis) or the husband is currently working in a low-wage job that he will subsequently quit (but that has a temporary income effect on female labor supply). In the latter case, the increase in the wife's labor supply would increase the probability of quitting because it acts as an indicator of his poor job match. On the other hand, if the expected increase in the husband's earnings associated with the quit is substantial and income can be smoothed over time, then the husband's decision to quit would induce a negative income effect on the wife's participation propensity. Thus, holding constant the wife's market wage rate, her reservation wage is a negative function of transactions costs associated with the quit and an uncertain function of the husband's expected wage gain. The wife's endogenous income share is

$$a_w^* = Z_3'P_3 + \gamma Q + \nu^*, \tag{9}$$

where a_w^* is the wife's desired income share, Z_3 contains the factors affecting the wife's market wage rate and her reservation wage rate, Q is the endogenous qualitative variable for the husband's quit with coefficient γ, and ν^* is distributed i.i.d. normally.

When the wife's market wage exceeds her reservation wage, the observed dependent variable in (9) is positive, but when her reservation wage is greater, the dependent variable forms a mass point at zero. This censoring of the dependent variable suggests the Tobit estimation of (9), specified as

$$a_w = \begin{cases} Z_3'P_3 + \gamma Q + \nu & \text{if } a_w^* > 0, \\ 0 \quad \text{if } a_w^* \leq 0, \end{cases} \tag{10}$$

where ν is an i.i.d. normal error term with mean zero. Joint consideration of (6) and (10) produces a probit-Tobit simultaneous equations model with the errors distributed bivariate normal with $E(\varepsilon_{it}, \nu_{it'}) = \sigma_{\nu\varepsilon}$, for t = t′, assuming that $\varepsilon_{it} \sim$ i.i.d. $N(0, \sigma_\varepsilon^2)$.

Another issue that is closely related to simultaneity is selectivity bias. Men with a propensity to quit may marry women who have preferences for market work or have the high relative market wage that would induce it. In each case, the coefficient on the share term would be erroneously positive: it would not reflect causality running from the wife's income share to the husband's quit. One way of reducing this erroneous correlation is to reestimate the quit equation using only the subsample of husbands whose wives work. This forces the quit to be a function of variation in the size of the wife's income share, not simply the wife's labor force status.

The conditional quit equation to be estimated is then:

$$Q_w = \begin{cases} 1 \quad \text{if } X_w'B_w > \varepsilon_w \\ 0 \quad \text{if } X_w'B_w \leq \varepsilon_w \end{cases} \text{if } a_w^* > 0, \tag{11}$$

where the qualitative quit decision for men with working wives, Q_w, is a function of X_w, which includes the wife's income share and has coefficients specific to the subsample with working wives. This equation is clearly subject to selectivity bias because of the conditioning on the wife working, $a_w^* > 0$. The technique used to correct for the selectivity bias is the standard Heckman (1979) two-stage method, which treats the selectivity problem like an omitted variable problem.[8] The wife's share is not endogenous as it is in the simultaneous equations estimates: in the case of selectivity bias the quit decision does not induce changes in the wife's income share.

Before turning to the estimation, it is helpful to restate the fundamental hypothesis of the paper.

PRIMARY HYPOTHESIS. The larger the wife's share of initial household income, the greater the propensity for the husband to quit, all else constant, because (1) the larger wife's share decreases the proportion of household wealth allocated to the risky quit option, thereby diversifying the allocation of wealth and increasing the expected value of quitting, and (2) the larger wife's share increases the liquidity constrained household's ability to finance transactions costs associated with quitting.

These relationships translate into a hypothesized positive coefficient on the wife's share of income in equation (6). Subsequent equations (8)–(11) introduce special cases of variation in this coefficient and its interpretation.

III. Empirical Results

The data set used to test the quit model is the Michigan PSID, a random survey of approximately 3,000 households interviewed annually from 1968 to 1981. More specifically, the subsample used here consists of married men who remained with their current employer or quit during the years 1971–77 and 1981. The years 1978–80 were dropped from the regressions because information on employer tenure is not available in those years. The years 1968–70 are dropped to limit the sample size used in the estimation of the nonlinear models and to reduce

[8] Assuming familiarity with the Heckman method, recall that the technique is consistent but not fully efficient since heteroscedasticity is introduced in the reformulated error term of the second stage. In the quit model above, the second stage estimation of the probit equation also relies on the assumption that the adjusted error term of the second stage is distributed normally—clearly a very strong assumption. Note, in addition, that eq. (11) could be one equation of a two-equation switching model, in which the other equation is for nonworking wives. However, the coefficients on the other independent variables are not of interest here, so the potential cross-correlation in the errors of the two regimes is not introduced.

Table 1
Mean Quit Rates by Age, Marital Status, and Wife's Labor Market Status

	18–24	25–34	35–44	45–54
Unmarried	.378	.188	.123	.074
	(725)	(944)	(389)	(245)
Married, wife not working	.232	.120	.050	.035
	(699)	(2,336)	(1,714)	(1,683)
Married, wife working	.230	.136	.075	.050
	(1,009)	(3,130)	(1,758)	(1,810)

NOTE.—Sample size is in parentheses (equal to the number of individuals times the numbers of years).

the number of survey years in which the tenure variable is categorical.[9] The work history is less than 8 years if the individual marries after 1971 or if data are missing for another reason.[10] Job change due to permanent layoff is excluded from the analysis in the year in which it occurs, and the age range is also limited to ages 16–59 to restrict quits to movement across employers. The resulting sample consists of 1,445 married males observed for approximately 8 years.

The quit rates for both married and unmarried men in the PSID are presented in table 1 by age group.[11] The measure of the quit rate based on this survey is biased downward because quit information is not obtained from a continuous time history. Instead, at each survey date the individual is asked the cause of the last termination of employment. Although this underestimates the quit rate, particularly for youths, striking regularities stand out. Quits do fall dramatically with marriage, across all age groups, but, on the other hand, the presence of a working wife generally offsets the decline in the quit rate with marriage. For prime-age married men (35–44), the quit rate is 50% greater for those with working wives than for those without working wives. For other age groups the differential is smaller, and it is nonexistent for the youngest group.

Given these interesting age patterns and the likely age variation in relative risk aversion, quit equations are estimated by age group. Four age groups are formed on the basis of 1971 age: ages 16–19, 20–29, 30–

[9] Until 1975, respondents were given the choice of seven discrete intervals for length of tenure. In constructing the TENURE variable (listed in table 2 below and discussed later in this paper), tenure is set equal to the means of the categorical intervals. After 1975 the variable is a continuous response.

[10] An individual is included in the data only during the years he is married. The results reported here are essentially unchanged when the sample is restricted to those who are married at least 1 year before and after the quit. Only the estimates of the quit equation for young men change modestly, with the wife's income variable becoming more significant.

[11] Note that these age groups differ from those presented elsewhere because these are grouped cross-sectionally, as compared to longitudinally.

39, and 40–49. Panel data impose considerable overlap in the resultant age groups across the 8 years. The actual age ranges across the panel years become 16–29, 20–39, 30–49, and 40–59, though it must be emphasized that much fewer men are distributed in the upper span of each age interval because there is a 3-year gap before the last year in the time series (1971–77 to 1981). Variable definitions and summary statistics by age group are provided in table 2.

Table 3 presents estimates of the standard quit equation. The equation is the probability of quitting as a function of the expected gain to quitting, $\bar{r}_q - r_s$, omitting risk aversion. The expected gain is a reduced-form function of education, experience, race, employer, tenure, the area unemployment rate, and the current wage rate.[12] Very generally speaking, the negative coefficients on experience, tenure, and education are likely to represent several simultaneous effects: job-specific skills may rise with tenure and schooling, thereby decreasing the relative value of alternative employment; tenure may be correlated with a high-quality job "match" with the current employer, thus implying a lower propensity to quit the job; experience decreases the time horizon for returns to new job-specific investment, though it may increase the quality of new job matches; and, last, all variables may be correlated with unobserved individual-specific preferences for mobility or unobserved ability. The current wage rate picks up many of these same effects.[13] However, a high wage more directly represents both a good "draw" from the wage offer distribution and a high-quality "match" with the current employer; thus higher wages clearly decrease the probability of quitting. No attempts have been made to place sufficient structure on the joint mobility–wage determination process so as to differentiate between the alternative effects listed above (for just such an analysis, see Topel [1986]).

[12] The expected gains to quitting are a function of the difference between the present values of income in the alternatives. Therefore it is likely that the reduced form is not linear in the X_1 variables of eq. (5a). However, the econometrician has so little information about expected alternative wage rates that linear reduced-form estimation is likely to be more informative than estimation as a function of imputed present values. Most micro quit equations take this approach (see Bartel 1979; Viscusi 1979*b*; and Topel 1986).

[13] Note that, for young men less than about age 25, the coefficient on the current wage has a strong positive effect rather than the hypothesized negative effect. This positive coefficient may indicate that high-wage youths have greater overall opportunities for wage growth than do low-wage youths, who could benefit more by staying with the current employer and earning returns to tenure. This pronounced age pattern would appear to warrant future investigation, but it is not the subject of this paper. There is also a possibility of simultaneity between wages and quit rates, but the reduced-form treatment of the expected alternative wage (or rate of return) makes this a possibility that is not readily addressed.

Table 2
Variable Definitions and Means (Married Men)
A. Definitions*

QUIT =	one for quit in last year
AGE =	age in survey year
BLACK =	one for race is black
EDUC =	years of education
CHILD =	number of children
WAGE =	natural log of the real wage, measured in cents
UNEMRT =	local area unemployment rate
TENURE =	years with current employer
EXPER =	years of labor market experience
WLABS =	ratio of wife's labor income to total household income
OTHS =	ratio of nonlabor income to total family income
TRANS =	ratio of transfer payments to total family income
EDWIFE =	years of education of wife
PUN =	predicted duration of unemployment (see text)
SAV =	one for savings greater than 2 months' income
MORT =	ratio of mortgage payment to household income

B. Means (by Age Group)†

Variables	16–29	20–39	30–49	40–59
QUIT	.215	.132	.071	.050
AGE	23.2	28.3	38.4	49.4
BLACK	.082	.077	.105	.088
EDUC	12.2	13.4	12.8	12.1
	(2.03)	(2.61)	(3.05)	(2.75)
CHILD	.80	1.37	2.39	1.42
	(.96)	(1.18)	(1.98)	(1.25)
WAGE	5.35	5.65	5.92	6.19
	(1.79)	(2.28)	(3.65)	(5.76)
UNEMRT	6.08	5.99	5.82	6.07
	(2.62)	(2.57)	(2.65)	(2.59)
TENURE	2.37	3.88	7.73	11.02
	(2.97)	(4.34)	(8.48)	(9.34)
WLABS	.194	.192	.142	.126
	(.21)	(.18)	(.17)	(.19)
OTHS	.008	.017	.022	.014
	(.05)	(.05)	(.09)	(.05)
TRANS	.038	.029	.028	.091
	(.13)	(.14)	(.14)	(.17)
EDWIFE	12.89	13.3	11.6	11.4
	(2.89)	(3.05)	(3.75)	(3.59)
EXPER	5.45	10.00	19.45	28.9
	(4.31)	(5.06)	(5.97)	(6.31)

* All variables are measured prior to the quit decision.
† Standard deviations are given in parentheses.

Note, however, that the tenure variable is more than simply a proxy for firm-specific human capital. Adding tenure to the quit equation is essentially equivalent to the insertion of a lagged dependent variable since it is the inverse outcome of the prior propensity to quit. Thus the implication of adding tenure can be easily described in the framework

Table 3
Standard Quit Equations (Married Men, by Age Group)

	16–29	20–39	30–49	40–59
EDUC	−.052	−.040	−.013	−.005
	(.031)	(.012)	(.016)	(.016)
TENURE	−.766	−.416	−.182	−.208
	(.015)	(.062)	(.042)	(.036)
EXPER	−.028	−.040	−.0037	.0007
	(.031)	(.009)	(.0074)	(.0072)
BLACK	.197	−.186	−.272	.090
	(.243)	(.098)	(.183)	(.170)
WAGE	.140	.009	−.103	−.195
	(.084)	(.036)	(.048)	(.058)
UNEMRT	.026	.018	−.007	−.019
	(.020)	(.012)	(.016)	(.021)
CONSTANT	.076	−.113	−.009	−.037
	(.163)	(.061)	(.109)	(.155)
Log likelihood	−366.42	−927.56	−478.98	−349.63
N	350	388	339	368
T	8	8	8	8

NOTE.—*N* is the number of individuals; *T* is the number of years. Standard errors are in parentheses. All equations contain the variables UNEMRT, OTHS, and TRANS.

of the renewal models of mobility between states. The insertion and signficance of tenure is a statement that the hazard rate is falling or that negative duration dependence exists. There is, of course, at least one clear reason for strong state dependence: longer duration in the state increases firm-specific human capital and thus lowers the conditional probability of future quits. As a lagged dependent variable, the tenure variable also picks up the mover-stayer effect, or worker heterogeneity in preferences for mobility that are unobserved by the econometrician. Thus adding the tenure variable should considerably lessen the effect that estimates of risk aversion have on the probability of quitting, assuming that preferences exhibiting risk aversion are highly correlated over time. In fact, the coefficient representing the effect of risk aversion, described below, is about 40% greater in size when tenure is omitted.

Turning to table 4, equation (6) is estimated by adding the wife's share of household income (measured prior to the quit decision) to the quit equation.[14] The effect of the wife's income share is significantly positive

[14] The estimated equations also contain two other measures of family assets: other family income (e.g., dividends and interest) and transfer payments (e.g., welfare), labeled OTHS and TRANS, respectively, when divided by total family income. These two variables have very different theoretical effects. Other income should have a positive asset effect on quits because, just like the wife's income, it is a less risky form of income and is uncorrelated with the husband's job choice. However, the effects of financial savings should be modeled in a context that permits the allocation of financial wealth across alternative financial assets

Table 4
Quit Equations (Married Men, by Age Group)

	16–29	20–39	30–49	40–59
WLABS	.605	.516	.529	.540
	(.322)	(.176)	(.247)	(.318)
CHILD	−.095	.062	−.040	−.025
	(.077)	(.041)	(.033)	(.022)
EDUC	−.072	−.057	−.009	−.018
	(.032)	(.014)	(.017)	(.017)
TENURE	−.764	−.411	−.176	−.198
	(.015)	(.061)	(.042)	(.033)
EXPER	−.033	−.048	−.006	−.008
	(.033)	(.010)	(.009)	(.007)
BLACK	.214	−.188	−.302	−.071
	(.255)	(.101)	(.186)	(.174)
WAGE	.155	.012	−.123	−.150
	(.083)	(.037)	(.055)	(.060)
CONSTANT	.264	−.027	−.318	.000
	(.163)	(.053)	(.109)	(.00)
Log likelihood	−362.3	−924.88	−475.35	−346.67
N	350	388	399	368
T	8	8	8	8

NOTE.—Standard errors are in parentheses. The equations also contain the variables UNEMRT, OTHS, and TRANS.

across all age groups, holding constant standard determinants of the quit propensity, including the husband's prior wage. Note that it is essential that the husband's wage rate be included; if excluded, the wife's income share would be driven in part by the husband's wage effect rather than her income effect.

If the wife's income variable is replaced by a dummy variable representing her labor force participation, the dummy variable is universally much weaker than the income share variable. Despite the obvious differences in mean quit rates by wife's labor force status (displayed in

---

as well as across alternative job decisions, thus considerably complicating the model. In any case, OTHS has the wrong sign for all age groups; it is negative and generally has an insignificant effect on quits. It seems very plausible that more risk-averse households may tend to save more, and thus the negative coefficient may be caused by the variable's positive correlation with risk-averse preferences that would decrease quits. One advantage of the wife's income variable is that it is very likely that risk aversion is not a primary determinant of the wife's participation decision. The second variable, transfer payments (TRANS), has a positive effect on quits that is sometimes significant. Transfer payments are likely to be correlated with poor skills and thus a higher quit rate—not a surprising effect, but one that is also outside the scope of this paper. Both OTHS and TRANS are likely to be susceptible to significant measurement error, and the mean values of these variables are very small relative to the wife's labor share (see table 2). The coefficient on the wife's share variable is not sensitive to the inclusion of these variables.

table 1), when other determinants of quitting are held constant, it is the wife's income that influences quits, not simply her market status. Even more remarkable, the magnitude of the effect of the wife's income share is very striking. Setting all independent variables equal to the mean values for the average married male, marriage to a woman who earns 40% of household income, compared to 0%, will increase the predicted quit rate from 4.4% to 6.4%, or by 45%, for prime-age males (30–49).

The results presented in table 4 assume that the errors are distributed i.i.d. normal. An alternative specification conforming to the less restricted covariance of equation (7) was also estimated. Random-effects estimation tends to result in higher-share coefficients. For example, the coefficient on WLABS for ages 30–49 rose to .729 (.325). The same occurred in tests for serially correlated errors, which displayed significant correlation, using a χ^2 test for its inclusion. Because of the relatively weak changes and the expense of estimating all the equations in the paper with a less restricted covariance matrix, only the results for the restricted covariance matrix are presented below.

A second specification test introduces the possibility of a wealth effect in the asset model: assuming decreasing relative risk aversion (rather than constant), an increase in total family wealth will increase the propensity to quit (see Sec. I). When the level of the wife's income or the level of family income replaces the share variable in the quit equation, the levels are positive, but not significant, across all age groups. In this simple test there is no evidence of a wealth effect, and I return to the assumption of constant relative risk aversion.

The results presented thus far are consistent with the implications of the mean-variance model but do not directly test the hypothesis that the coefficient on the wife's income share represents constant relative risk aversion times the variance of wages. These two later variables are now assumed to be a function of a set of observables, defined as $Z_1 + Z_2$ in equations (8a) and (8b), including education, age, race, family size, and the local unemployment rate.[15] There is, unfortunately, very little evidence regarding the likely signs of these variables in the equations for risk aversion and the variance of the alternative wage distribution. Previous empirical research that tests the constancy of relative risk aversion generally does not have the data necessary for tests of variation of ρ across demographic groups.[16] In the process of estimating a consumption

[15] The PSID contains a variable labeled "risk preference," an ordinal measure increasing in magnitude when the individual does not smoke and has substantial insurance, a late model car, and substantial savings. Since there are no controls for causal determinants of these variables, the risk variable is not utilized here.

[16] See Friend and Blume (1975). Note also that the literature on redlining and borrowing constraints has not been very successful in separating demand and supply-side factors necessary in isolating demographic correlates with risk aversion (see Benston and Horsky 1979; and Avery 1981).

function subject to liquidity constraints, Mariger (1986) estimates education interactions with rates of time preference, which are analogous to risk aversion. He finds a statistically significant decline in constant relative risk aversion with increasing education. The organizational behavior literature has also looked at concepts that they believe are similar to risk aversion, one being "need achievement," defined to be the motivation for achievement moderated by the fear of failure. This variable rises with education and declines with age in experimental studies (e.g., Veroff et al. 1974). Thus these correlations are consistent with survey results regarding portfolio selection (Cohn et al. 1975) and with one's general perceptions: risk aversion rises with age and declines with education. No such sign predictions, however loose the bounds, can be made about the variance of the alternative wage distribution, σ_q^2. There is evidence regarding the unconditional variance of wages by age, education, and occupation, but these results may not be representative of the likely variance of the individual's true alternative wage distribution.[17] Last, demographic variation in the coefficient on the wife's labor share may also pick up variation in the transactions costs, as specified in equation (5b). More explicit variation in transactions costs is considered subsequently.

Table 5 presents the estimates of these numerous interactions with WLABS. Across all age groups, the addition of the set of interactions is significant above the 5% level for the likelihood ratio test. The absolute values of these coefficients cannot be interpreted in a probit equation, so table 6 presents several predicted probabilities of quitting, which vary with the wife's labor income share and the husband's education and race, conditional on the sample mean values for all other variables.[18] The implications are as follows. (1) For white males, the coefficient on the wife's share declines with education and at a decreasing rate. This is consistent with the above evidence that constant relative risk aversion decreases with education, but it may also represent decreasing transactions costs. (2) The effects are quite different for blacks than for whites. The generally negative coefficient on the black constant term in table 4 indicates that blacks tend to have lower quit rates than whites. The interactions estimated in table 5 demonstrate that only the more educated blacks are more sensitive to the wife's share of income—exactly the opposite of the effect for whites. Since the transactions costs associated with quitting tend to fall with education (i.e., unemployment falls with

[17] For example, Parsons (1978) shows that the variance rises with education, and King (1974) shows it is greater for technical occupations. The individual's true alternative wage distribution need not be consistent with these findings because there are virtually no controls for individual ability or experience.

[18] The mean values of right-hand-side variables, such as the wage rate, are fixed across education and race groups so that the predicted probabilities measure only variation in the coefficients.

Table 5
Quit Equations with WLABS Interactions (Married Men, by Age Group)

	16–29	20–39	30–49	40–59
WLABS	20.06	5.63	3.70	3.15
	(5.44)	(3.84)	(2.01)	(2.11)
BLACK	−.385	−.293	−.287	−.063
	(.267)	(.251)	(.244)	(.154)
BLACK · WLABS	−19.44	−13.10	−12.27	−10.01
	(15.20)	(6.05)	(4.14)	(6.35)
BLACK · EDUC · WLABS	1.53	.831	.926	1.36
	(1.26)	(.452)	(.318)	(.41)
EDUC	.014	.084	.021	.083
	(.012)	(.057)	(.017)	(.063)
$EDUC^2$	−.001	−.006	.001	−.004
	(.0006)	(.001)	(.001)	(.002)
EDUC · WLABS	−2.56	−.72	−.368	−.240
	(.79)	(.35)	(.171)	(.196)
$EDUC^2$ · WLABS	.085	.025	.015	.014
	(.030)	(.012)	(.008)	(.008)
UNEMRT · WLABS	.126	.063	.055	.058
	(.070)	(.051)	(.059)	(.050)
AGE · WLABS	−.073	−.016	−.036	−.031
	(.100)	(.037)	(.042)	(.036)
TENURE	−.691	−.131	−.043	−.049
	(.082)	(.023)	(.011)	(.007)
EXPER	−.025	−.044	−.007	−.008
	(.026)	(.017)	(.008)	(.008)
WAGE	.110	−.036	−.159	−.220
	(.030)	(.047)	(.049)	(.061)
CONSTANT	.082	−.090	.0001	.000
	(.135)	(.008)	(.0000)	(.000)
Log likelihood	−345.2	−899.6	−461.2	−330.6
N	350	388	339	368
T	8	8	8	8

NOTE.—Standard errors are in parentheses. The equations also contain the variables UNEMRT, OTHS, and TRANS.

education), the implication is that the more educated blacks are either more risk averse or face a much higher variance of alternative wages.

While these results appear to be consistent with individual variation in constant relative risk aversion, the researcher's inability to form

Table 6
Predicted Quit Probabilities (Married Men, Aged 30–49)

	White	Black
EDUC, WLABS:		
12, 0	13.80	8.40
12, .40	20.07	2.10
16, 0	7.66	4.29
16, .40	10.60	15.20

Table 7
Quit Equations with Liquidity Variables (Married Men)

	2*	3†	4‡
WLABS	.469 (.224)	.507 (.199)	.566 (.219)
PUN	−.015 (.033)	...	...
PUN · WLABS	.24 (.79)	...	...
SAV	...	−.20 (.11)	...
SAV · WLABS	...	−.073 (.407)	...
MORT	...	...	1.37 (1.93)
MORT · WLABS	...	...	−.26 (.33)

NOTE.—Standard errors are in parentheses. The probits also contain the variables CHILD, EDUC, TENURE, EXPER, BLACK, WAGE, UNEMRT, OTHS, and TRANS.
* Ages 20–49 pooled; all years; $N \cdot T$ = 5,820.
† Only years 1971, 1972, 1975; all ages; $N \cdot T$ = 2,820.
‡ Excludes years 1973, 1974, 1975; all ages; $N \cdot T$ = 7,225.

hypotheses regarding the distribution of alternative wages hampers any clear conclusion. Additional evidence is also needed to separate risk aversion and transactions costs.

One important transactions cost associated with quitting is the ex ante expectation of unemployment. Predicted unemployment duration conditional on quitting, PUN, is calculated from a regression of weeks of unemployment for quitters as a function of all X variables.[19] The expected unemployment duration is subject to selection bias, but, assuming that the predicted duration is biased downward for those who do not quit, the selection bias would be in the direction of not rejecting the transactions cost hypothesis. In table 7, the interaction between PUN and WLABS is insignificantly different from zero. In addition, the likelihood ratio test rejects the addition of the predicted unemployment

[19] Two other unemployment effects are tested. One is the interaction of the local area unemployment rate with the wife's labor share, an interaction that is never significant. In the second test, the actual levels of unemployment associated with quitting are examined. Only 38% of all quitters experience unemployment sometime during the year in which they quit (either before or after the quit), and it is of short mean duration—1.4 weeks with a standard deviation of 2 weeks (rising with age). Therefore the quit equation is estimated only for those individuals who experienced no unemployment. The coefficient on wife's labor share, at .479 for pooled ages 20–49, does not decline significantly. Of course, the equation for those continuously employed is biased by selectivity and considers only the ex post quit experience, not ex ante expectations.

variable and its interaction with WLABS. Given the short mean duration of unemployment, at 1.4 weeks for those quitters who experience unemployment, it is not too surprising that expected unemployment does not appear to play a very significant role in explaining the wife's income effect on the quit decision.

The PSID data set contains two variables that may act as proxies for the liquidity constraints faced by the husband in making the quit decision. The first variable is a yes/no response to the question, "Do you currently have more than two months of income in savings (in checking, savings, bonds, etc.)?" The question is asked in only 3 years of the survey, when approximately 78% of the sample answer yes. This dummy variable is useful here because it is highly unlikely that a spell of quit-unemployment will last more than 2 months. Column 2 of table 7 shows that the presence of savings decreases the propensity to quit but that the savings dummy variable has no significant effect on the coefficient on the wife's income share. As a proxy for the reduction of liquidity constraints, the savings dummy should increase quits, but it is also quite likely that households with high savings are more risk averse and thus less likely to quit.[20] The second variable is the value of home mortgage payments; higher mortgage payments may deter quits because of the fear of default. The mortgage variable used in table 7, the percent of income spent on mortgage payments, has a weak positive effect on quits (possibly a wealth effect) but has no effect on the coefficient on the wife's income share.

Researchers have done very little empirical work examining the economic effects of liquidity constraints and risk aversion because it is inherently difficult to separate the two effects (King 1974). However, the insignificance of the savings and mortgage data seems to suggest that, while transactions costs may play a role in the quit decision, they do not seem to be a dominant explanation for the positive effect of the wife's share of income on male quits.

Two final tests of the model incorporate simultaneity and selectivity bias, as discussed in Section I above. Table 8 contains the estimated coefficients for the system of simultaneous equations specified in equations (6) and (10). The estimation is maximum likelihood for a dummy endogenous censoring model with cross-equation correlation of errors (probit-Tobit framework).[21] Across all age groups, simultaneity does

[20] The savings variable may also pick up a positive asset effect that is analogous to the effect of the wife's income share on the husband's quit propensity. However, the effect of the savings variable has some important conceptual differences from that of the wife's income (for a discussion of this point, see n. 14 above).

[21] The estimation is done using the HOTZTRAN program by Robert Avery and V. Joseph Hotz, Carnegie-Mellon University. Estimates of the correct

Table 8
Simultaneous Equations Estimation (Married Men, by Age Group)
A. Quit Equation

	16–29	20–39	30–49	40–59
WLABS	.532	.503	.516	.507
	(.320)	(.184)	(.251)	(.311)
CHILD	−.077	.071	−.034	−.011
	(.075)	(.063)	(.032)	(.017)
EDUC	−.068	−.060	−.008	−.017
	(.032)	(.013)	(.017)	(.017)
TENURE	−.803	−.494	−.184	−.188
	(.013)	(.071)	(.042)	(.034)
EXPER	−.030	−.035	−.004	−.008
	(.031)	(.011)	(.009)	(.008)
BLACK	.070	−.211	−.319	−.054
	(.242)	(.124)	(.190)	(.165)
WAGE	.159	−.073	−.110	−.157
	(.085)	(.059)	(.056)	(.058)

B. WLABS Equation

	16–29	20–39	30–49	40–59
QUIT	.014	.072	.086	.066
	(.017)	(.037)	(.026)	(.043)
CHILD	−.093	−.063	−.055	−.020
	(.017)	(.015)	(.008)	(.022)
EDUC	−.004	−.009	−.022	−.038
	(.007)	(.009)	(.004)	(.014)
EDWIFE	.014	.019	.015	.009
	(.006)	(.006)	(.005)	(.008)
AGE	.004	.004	.007	−.001
	(.003)	(.003)	(.001)	(.001)
BLACK	.003	.051	.042	.040
	(.046)	(.031)	(.036)	(.021)

NOTE.—Standard errors are in parentheses. Quit equation contains variables UNEMRT, OTHS and TRANS.

appear to characterize the decision: the share equation appears to be well specified, and the cross-equation residuals are correlated.[22] The weak negative effect that CHILD has on the husband's quit rate appears to operate largely through the effect that it has on reducing the wife's share of income. However, there is no real evidence of simultaneous equations bias in the WLABS coefficient, except for the youngest age group, whose WLABS coefficient rises.[23] Thus, while the positive coef-

asymptotic errors are presented. Two-stage methods of estimation of the simultaneous probit-Tobit model are also available (see Maddala 1983, p. 246).

[22] For example, for ages 30–49, $\sigma_{v\varepsilon} = .21$ (.01).

[23] Simultaneity between the husband's quit decision and the wife's income share is most likely to be marked by an increase in her share before his quit. To test this prediction, a new variable is constructed that is the average of the wife's

ficient on wife's income share may well reflect the hypothesis that a greater male propensity to quit causes female labor supply to rise, the data and techniques at hand find little evidence of this causal relation.

Similarly, no evidence of selectivity bias is found. To implement Heckman's method (discussed above), the probit equation for the wife working is estimated as a function of the variables in the WLABS equation of table 8 and is used to calculate λ, which is inserted in the quit equation that is conditional on the wife working (men without working wives are dropped). In three of the four age groups the signs on the λ-variable are in the predicted positive direction (these results are not shown). However, likelihood ratio tests reject the addition of the λ in all age groups. The coefficient values for WLABS are not directly comparable to those in table 4 because the true coefficients are now normalized by the variance of the adjusted residual. Nevertheless, the two WLABS coefficients remain strong, indicating that it is the size of the wife's labor share that matters, not simply whether she works. Thus these tests for selectivity and simultaneity bias have not diminished the conclusion that an increase in the wife's share of income appears to cause an increase in the husband's quit rate.

In conforming to the model of Section I, the empirical analysis has been confined to married men. Yet, before concluding, it is interesting to extend this analysis to estimate the total effect that martial status has on the male quit propensity. Looking at these results, presented in table 9, the marriage dummy variable has its standard negative effect on the quit propensity of men. However, the marriage dummy does not represent the complete effect that marriage has on the quit rate. As estimated in the adjacent columns, an increase in the wife's labor income share can significantly offset the negative effect that marriage has on quits. The best way of assessing the magnitude of the offsetting effect is to calculate the predicted probabilities of quitting across alternative values of the wife's income. The predicted probabilities of quitting evaluated at the mean values of the independent variables of the married subsample of men aged 30–49 (with children set equal to zero) are as follows: married with wife working (WLABS = .40) results in a predicted

income share over the entire longitudinal history of the couple. The probit coefficients on this average share variable are very similar to those on WLABS in table 4; they are .52 (.37), .58 (.24), .69 (.31), and .56 (.43) for the four age groups, respectively (standard errors are in parentheses; all other control variables are the same as those in table 4). There is no indication that quits are dramatically influenced by temporal shifts in the share variable. Furthermore, the coefficient on the average share tends to rise with age. These are both characteristics of the effects of risk aversion, as compared to the need to finance expected unemployment costs.

Table 9
Quit Equations (Married and Unmarried Men, by Age Group)

	16–29		20–39		30–49		40–59	
MARRY	−.176	−.246	−.053	−.140	−.149	−.229	−.288	−.385
	(.104)	(.128)	(.092)	(.100)	(.122)	(.130)	(.177)	(.186)
WLABS	. . .	.421	. . .	.408	. . .	.615	. . .	.641
		(.318)		(.176)		(.249)		(.315)
CHILD	−.057	−.054	.003	.013	−.058	−.050	−.023	−.015
	(.065)	(.068)	(.038)	(.039)	(.029)	(.031)	(.035)	(.034)
EDUC	−.018	−.046	−.026	−.044	−.003	−.009	−.010	−.021
	(.021)	(.020)	(.012)	(.012)	(.014)	(.015)	(.015)	(.015)
TENURE	−.710	−.700	−.424	−.410	−.187	−.182	−.198	−.192
	(.014)	(.012)	(.047)	(.057)	(.038)	(.039)	(.033)	(.033)
EXPER	−.032	−.039	−.025	−.031	−.003	−.009	−.002	−.007
	(.021)	(.015)	(.011)	(.011)	(.006)	(.007)	(.007)	(.007)
BLACK	.025	.031	−.212	−.228	−.229	−.247	−.048	−.069
	(.171)	(.161)	(.096)	(.098)	(.166)	(.167)	(.185)	(.187)
WAGE	.084	.152	−.043	−.003	−.088	−.059	−.139	−.100
	(.058)	(.032)	(.034)	(.036)	(.043)	(.047)	(.055)	(.055)
CONSTANT	.115	.00	−.092	.00	−.434	−.409	−.123	−.040
	(.151)	(.00)	(.064)	(.00)	(.102)	(.099)	(.152)	(.155)
Log likelihood	−549.34	−543.4	−1137.4	−1131.7	−570.68	−566.4	−410.1	−406.9
N	460	460	540	540	358	358	387	387
T	8	8	8	8	8	8	8	8

NOTE.—Standard errors are in parentheses. The equations also contain variables UNEMRT, OTHS, and TRANS.

probability of 6.4%; married with wife not working, 4.4%; and unmarried, 6.4%. Assuming equal mean values of the independent variables, the effect of a working wife's income can completely offset the negative marital effect on quits. However, there is a greater difference between married and unmarried quit rates, resulting from the differences in the values of the independent variables. When the probability of quitting for unmarried men is calculated at the mean values of *their* independent variables, the probability rises to 7.1%. They quit more because they have lower levels of wages and tenure than married respondents, holding experience constant.[24]

IV. Conclusions

Married men with working wives seem to have considerably higher quit rates than those without. Using the PSID data, the empirical results indicate that an increase in the wife's share of family income does have a very strong positive effect on the quit propensity. The predicted probability of quitting for prime-age married males rises from 4.4% to 6.4% when the wife's share of income increases from 0% to 40%.[25] The cause of this appears to be the role that the wife's income plays in stabilizing family income while the husband takes on the risky option of quitting. No empirical support is found for the alternative hypothesis, that a greater male quit propensity causes the wife's labor supply to rise, though this alternative may be a underlying factor. One simple conclusion appears to justified: risk-averse husbands who are the sole support of their families are less likely to quit their jobs and risk the potential income losses. While many researchers have emphasized the likely importance of the effects of risk aversion on human capital investment decisions (Weiss 1972; and Levhari and Weiss 1974), these estimates of the quit propensity find potential empirical support for its significance.

The strength of the positive effect of the wife's income on male quit rates has several important implications. First, the tremendous increase in the labor force participation rate of married women in the last 2 decades may well account for a noticeable increase in the male quit rate. There arose speculation in the 1960s that a "new industrial feudalism" may be causing quit rates to decline because of an increase in firm-specific human capital and institutions, such as pensions, that result in greater attachment to the firm. Time-series data are available only for

[24] Note that the means of the independent variables for married men do not vary with the wife's labor force status.

[25] For younger age groups, the level of the quit rate is higher and the percent gap between predicted quit rates for wife-working and nonworking households is slightly smaller. Of course, these point estimates depend on the specification of the equation. When interactions with WLABS are added, the gap is somewhat smaller, and when the residual covariance is less restricted, the gap is larger.

manufacturing, where it is not surprising to find a recent decline in quit rates as unionized manufacturing workers age (Ragan 1984). For the economy as a whole, it is possible that the dramatic increase in female participation may be counteracting any hypothesized downward trend. Extrapolating our data to time-series changes in average quit rates, the historical increase in female participation rates would have caused male quit rates to rise by approximately 20% for prime-age males over the last 2 decades.

Second, a rising quit rate implies an increase in frictional unemployment, and thus a growing natural rate, but these trends would reflect improvements in market efficiency. Quitting is by far the dominant form of employee turnover, relative to layoffs and firings, particularly for younger men.[26] Young married women have experienced the greatest increase in their labor force participation rate, rising from 29% to 63% over 1960–83 for those aged 25–34. This increase in their participation is particularly important in analyzing male quit rates because marriage has the strongest negative effect on quits during the earlier periods of the male life-cycle. This is also the period during which job matching and job search models would indicate the greatest efficiency gains from job search. Thus, while increasing the level of frictional unemployment, an increase in the wife's income share also increases the likelihood that the husband will choose the job that maximizes his expected monetary return or best utilizes his skills.

References

Avery, Robert. "Estimating Credit Constraints by Switching Regressions." In *Structural Analysis of Discrete Data with Econometric Applications*, edited by C. Manski and D. McFadden. Cambridge, Mass.: MIT Press, 1981.

Bartel, Ann. "The Migration Decision: What Role Does Job Mobility Play." *American Economic Review* 69 (December 1979): 775–86.

Benston, George J., and Horsky, Dan. "Redlining and the Demand for Mortgages in the Central City and Suburbs." *Journal of Bank Research* 10 (Summer 1979): 72–87.

Cohn, Richard; Lewellen, Wilbur G.; Lease, Ronald C.; and Schlarbaum, Gary G. "Individual Investor Risk Aversion and Investment Portfolio Composition." *Journal of Finance* 30 (May 1975): 605–20.

Freeman, Richard. "The Effect of Unions on Worker Attachment to Firms." Working Paper no. 400. Cambridge, Mass.: National Bureau of Economic Research, October 1979.

[26] The annual individual quit rate in the last 15 years averaged 7.8%, compared to 2.9% for all employer-initiated separations (Freeman 1979, based on household surveys).

Friend, Irwin, and Blume, Marshal. "The Demand for Risky Assets." *American Economic Review* 65 (December 1975): 900–922.
Hause, John. "The Risk Element in Occupational and Educational Choices: A Comment." *Journal of Political Economy* 82 (July 1974): 803–7.
Heckman, James. "Sample Selection Bias as a Specification Error." *Econometrica* 47 (January 1979): 153–63.
Johnson, William. "A Theory of Job Shopping." *Quarterly Journal of Economics* 92 (May 1978): 261–78.
King, A. "Occupational Choice, Risk Aversion, and Wealth." *Industrial and Labor Relations Review* 27 (July 1974): 506–96.
Levhari, David, and Weiss, Yoram. "The Effect of Risk Aversion on Investment in Human Capital." *American Economic Review* 64 (December 1974): 950–63.
Lippman, Steven, and McCall, John J. "The Economics of Job Search: A Survey." *Economic Inquiry* 19 (May 1976): 155–89.
Maddala, G. S. *Limited-dependent and Qualitative Variables in Econometrics.* Cambridge: Cambridge University Press, 1983.
Mariger, Randall P. *Consumer Behavior and the Effects of Government Fiscal Policies.* Cambridge, Mass.: Harvard University Press, 1986.
Parsons, Donald O. "The Autocorrelation of Earnings, Human Wealth Inequality, and Income Contingent Loans." *Quarterly Journal of Economics* 92 (November 1978): 551–69.
Ragan, James F., Jr. "Investigating the Decline in Manufacturing Quit Rates." *Journal of Human Resources* 19 (Winter 1984): 53–71.
Samuelson, Paul. "The Fundamental Approximation Theorem of Portfolio Analysis in Terms of Means, Variances and Higher Moments." *Review of Economic Studies* 37 (October 1970): 537–42.
Topel, Robert H. "Job Mobility, Search, and Earnings Growth: A Reinterpretation of Human Capital Earnings Functions." In *Research in Labor Economics,* edited by Ronald Ehrenberg. Greenwich, Conn.: JAI, 1986.
Veroff, J.; Atkinson, J. W.; Feld, S. C.; and Gurin, G. "Use of Thematic Apperception to Assess Motivation in a Nationwide Interview Study." In *Motivation and Achievement,* edited by John Atkinson and Joel Raynor. New York: Wiley, 1974.
Viscusi, W. Kip. "Job Hazards and Worker Quit Rates: An Analysis of Adaptive Worker Behavior." *International Economic Review* 20 (February 1979): 29–58. (*a*)
———. "Sex Differences in Working Quitting." *Review of Economics and Statistics* 61 (1979): 388–97 (*b*)
Weiss, Yoram. "The Risk Element in Occupational and Educational Choice." *Journal of Political Economy* 80 (November/December 1972): 1203–13.
———. "The Wealth Effect in Occupational Choice." *International Economic Review* 17 (June 1976): 292–307.