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TURNOVER AND THE DYNAMICS OF LABOR DEMAND

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ABSTRACT

The theory of the dynamics of labor demand is based either on the costs of adjusting the level of employment or on the costs of hiring or firing (of gross changes in employment). We write down a generalized cost of adjustment function that includes both types of cost and allows for asymmetries in those costs. We derive the firm's rational-expectations profit - maximizing path of employment demand and the Euler equation whose parameters we estimate.

Identifying the two types of costs requires complete data on turnover, which were available for the U.S. through 1981. We use these data for manufacturing to demonstrate that both types of adjustment cost figure in the representative firm's profit-maximizing decisions about employment, and that both types of cost are asymmetric (leading here to quicker increases than decreases in employment).

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I. Introduction

Since the early 1980s research on the dynamic demand for labor has been burgeoned. This literature has been based in some cases on the costs of changing the level of employment (e.g., Shapiro, 1986), in others on the costs of hiring and firing (e.g., Nickell, 1986). All the estimation, though, has been based on net changes in employment (on variations in employment levels).

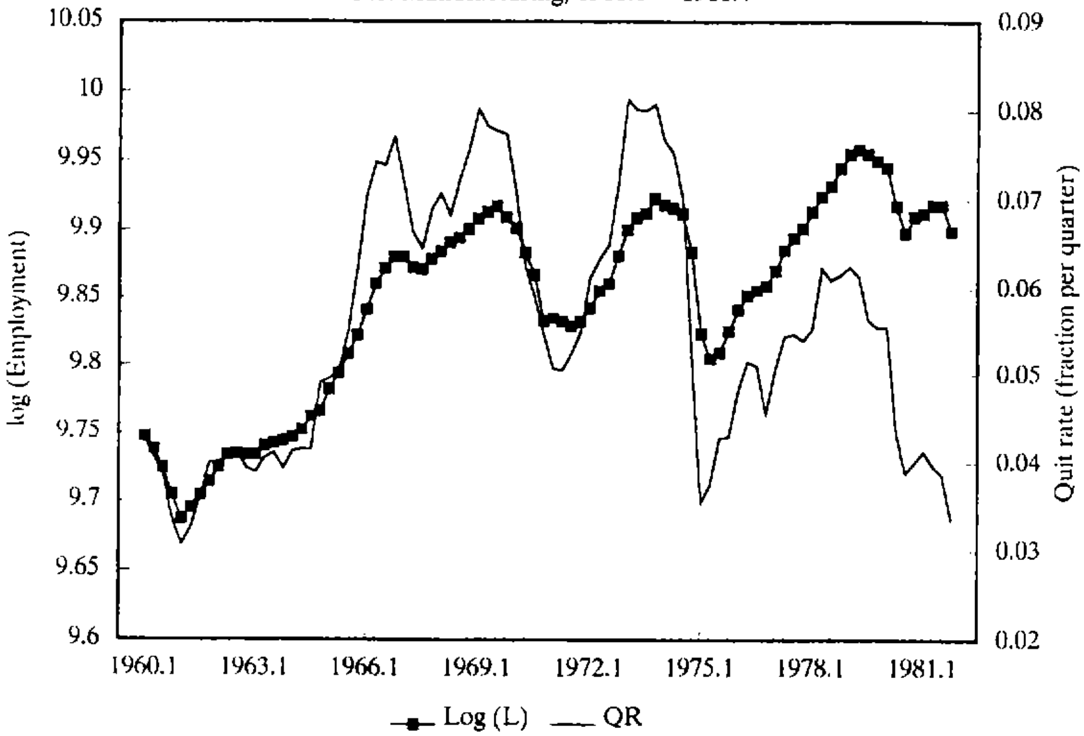
Here we examine models of dynamic labor demand by accounting for variations in voluntary mobility (job–quitting) that are linked to net changes in employment, ΔL, by:

(1) $\Delta L \equiv H - [F + Q],$

where H, F and Q are the numbers of workers per period who are hired, fired (laid off) and who quit. We assume that $F \bullet H \equiv 0$, i.e., concurrent hiring and firing of homogeneous workers does not take place. We also assume that employers choose the number of workers to hire or fire based upon a quit rate that they see as exogenous.[1] That this extension of the literature on labor–market dynamics may be important is demonstrated by Figure 1, which shows for manufacturing in the United States for 1960–1981 the quarterly quit rate along with the the logarithm of the level of employment. The rate of quitting is huge; and the quit rate is highly cyclical and closely correlated with employment.[2] Because quits drop precipitously during slumps, it is much harder for employers to cut their work forces then.

Figure 1. Log (Employment) and Quit Rate

U.S. Manufacturing, 1960:1 – 1981.4



Given the variability of quits, including them in dynamic econometric models of labor demand is important for two reasons. First, the theoretical models include the user cost of labor, which depends on the quit rate. Thus excluding quits from the estimation yields biased estimates of all the other parameters, especially given the size and variability of the quit rate. A few other studies have been concerned about this (e.g. Burgess and Nickell, 1990; Lockwood and Manning, 1992), but none has included a time series on the quit rate in the estimation.[3] Second, the growing literature on business-cycle asymmetry arising from differences in the costs of hiring and firing (Burgess, 1991; Pfann and Palm, 1992) has been unable to link the hypothesized asymmetry to the underlying hiring/firing costs except under the obviously incorrect assumption that the quit rate is constant. The estimates here thus provide a closer link to the underlying theory of asymmetric adjustment costs.

The failure of studies of the dynamic demand for labor to account for these large flows is no doubt due to the absence of time series on voluntary turnover in the major industrialized countries over which these studies have been estimated. Yet data on quit rates were collected on a monthly basis for manufacturing in the United States, though only through 1981; and the effects of quits on labor-market outcomes, including employment demand, was frequently studied in the 1960s and 1970s (e.g., Behman, 1964; Brechling, 1975). Since much of the recent literature has been based on sophisticated econometric methods for examining macroeconomic dynamics, we use these new methods, in particular, generalized method of moments (GMM) estimation (Hansen, 1982). Our results can thus be viewed as resurrecting an important labor–market measure, the rate of voluntary turnover, in recognition of its role in the modern theory of dynamic factor demand and in the estimation of models based on that theory.

II. An Estimable Model of Labor Demand in the Presence of Quits

We assume that the representative firm faces a quadratic production function:

$$(2) \qquad Y_t = (\alpha + \epsilon_t)X_t - .5X_t'AX_t \text{ , where } X_t = (L_t, K_t) \text{ ,}$$

where $\epsilon_t = (\epsilon_{1t} , \epsilon_{2t})'$ is a vector of disturbances reflecting the impact of random shocks on the optimization process; α is a 2x1 vector of parameters, and A is a 2x2 positive–definite matrix of parameters. Costs consist of the variable static costs of labor:

$$VLC(L_t, W_t) = W_tL_t \text{ ,}$$

and the costs of adjusting the demand for workers:

$$(3) \qquad AC = AC(\Delta L_t, Q_t, \theta) \text{ ,}$$

where θ is a vector of parameters. The specification of (3) distinguishes among the models estimated here. However, because (3) includes Q_t this general class of models differs from and expands on those estimated elsewhere.

All three models that we estimate specify versions of (3) allowing for quadratic adjustment costs in both the net change in employment and the difference between the endogenous hires and separations. Model I specifies AC as:

$$\text{(4a)} \qquad AC(\Delta L_t, Q_t, \theta) = .5\theta_2(\Delta L_t)^2 + .5\theta_3(\Delta L_t + Q_t)^2 ,$$

with $\theta_2, \theta_3 \geq 0$. The terms in θ_2 and θ_3 embody the possibility that both net and gross changes in employment generate adjustment costs, as in Hamermesh (1992) in the micro context. In the context of the representative firm the two terms can be interpreted as implying that there are increasing marginal costs of changing employment (hiring or laying off) and that these costs are greater when more workers have quit.

Model II allows for the additional possibility of asymmetric adjustment depending on whether ΔL_t is positive or negative:

$$\text{(4b)} \qquad AC(\Delta L_t, Q_t, \theta) = -1 - \theta_{11}\Delta L_t + \exp(\theta_{11}\Delta L_t) + .5\theta_2(\Delta L_t)^2 + .5\theta_3(\Delta L_t + Q_t)^2 .$$

This asymmetry is specified using the functional form proposed by Pfann and Palm (1992) that was very useful in describing British and Dutch time series on manufacturing employment.

A more general model allows for quadratic adjustment costs on both net changes in employment and on hires/layoffs, as in Model I. It allows for asymmetric responses to positive and negative net changes in employment, as in Model II; and it also specifies the possibility that there is asymmetric adjustment depending on whether $\Delta L_t + Q_t \equiv H_t - F_t$ is positive or negative. In this Model (III) adjustment costs are:

(4c) $$AC(\Delta L_t, Q_t, \theta) = -2 - \theta_{11}(\Delta L_t) + \exp(\theta_{11}\Delta L_t) - \theta_{12}(\Delta L_t + Q_t) + \exp(\theta_{12}(\Delta L_t + Q_t)) + .5\theta_2(\Delta L_t)^2 + .5\theta_3(\Delta L_t + Q_t)^2 .$$

Even more general specifications could be written down, e.g., allowing for interactions between ΔL_t and Q_t; but (4c) is the simplest general specification that allows for both gross and net adjustment costs and the possibility of asymmetric adjustment by a representative firm employing homogeneous labor.

The firm maximizes the objective function:

$$\underset{L_t}{Max} E_t[\sum_{i=0}^{\infty} \beta^i [Y_{t+i} - VLC_{t+i} - AC_{t+i}]] \quad ,$$

with respect to L_t, where E_t is the expectations operator conditional on the information available at time t, and $\beta < 1$ is the discount factor. We assume that decisions about L occur simultaneously with the flow of quits, Q. Assuming adjustment costs are described by (4c), for given values of K_t, Q_t and ϵ_t the representative firm operates each period according to the Euler equation:

(5) $$\alpha_1 + \epsilon_{1t} - \alpha_{11}L_t - \alpha_{12}K_t - W_t + \theta_{11} - \theta_{11}\exp(\theta_{11}\Delta L_t) + \theta_{12} - \theta_{12}\exp(\theta_{12}(\Delta L_t + Q_t)) - \theta_2\Delta L_t - \theta_3(\Delta L_t + Q_t) + E_t\{\beta[-\theta_{11} + \theta_{11}\exp(\theta_{11}\Delta L_{t+1}) - \theta_{12} + \theta_{12}\exp(\theta_{12}(\Delta L_{t+1} + Q_{t+1})) + \theta_2\Delta L_{t+1} + \theta_3(\Delta L_{t+1} + Q_{t+1})]\} = 0 .$$

Let $\widetilde{\theta}_2 = \theta_2 + \theta_3$, and $\widetilde{\alpha}_1 = (\theta_{11} + \theta_{12})(\beta - 1) - \alpha_1$. The values of the variables realized during period t+1 are substituted for the unobserved one-period ahead expectations of employment and quits, and a forecast error η_{t+1} is appended. Then (5) becomes:

$$(6) \qquad \beta\Delta L_{t+1} - \Delta L_t = \tilde{\alpha}_1/\tilde{\theta}_2 + (\alpha_{11}/\tilde{\theta}_2)L_t + (\alpha_{12}/\tilde{\theta}_2)K_t + (1/\tilde{\theta}_2)W_t + (\theta_3/\tilde{\theta}_2)[-\beta Q_{t+1} + Q_t] + (1/\tilde{\theta}_2)\{\theta_{11}[-\beta\exp(\theta_{11}\Delta L_{t+1}) + \exp(\theta_{11}\Delta L_t)] + \theta_{12}[-\beta\exp(\theta_{12}(\Delta L_{t+1} + Q_{t+1})) + \exp(\theta_{12}(\Delta L_t + Q_t))]\} - \epsilon_{1t}/\tilde{\theta}_2 + \eta_{t+1} .$$

The estimating equation is:

$$(7) \qquad L_{t+1} = \overline{\alpha}_1 + \overline{\alpha}_{11}L_t - \beta^{-1}L_{t-1} + \overline{\alpha}_{12}K_t + \overline{\theta}_2 W_t + \overline{\theta}_3\overline{\Delta}Q_t + \overline{\theta}_2\{\theta_{11}[-\beta\exp(\theta_{11}\Delta L_{t+1}) + \exp(\theta_{11}\Delta L_t)] + \theta_{12}[-\beta\exp(\theta_{12}(\Delta L_{t+1} + Q_{t+1})) + \exp(\theta_{12}(\Delta L_t + Q_t))]\} + \tilde{\eta}_{t+1},$$

where $\overline{\alpha}_1 = \tilde{\alpha}_1(\beta\tilde{\theta}_2)^{-1}$; $\overline{\alpha}_{11} = ((1 + \beta)\tilde{\theta}_2 + \alpha_{11})(\beta\tilde{\theta}_2)^{-1}$; $\overline{\alpha}_{12} = \alpha_{12}(\beta\tilde{\theta}_2)^{-1}$; $\overline{\theta}_2 = (\beta\tilde{\theta}_2)^{-1}$; $\overline{\theta}_3 = \theta_3(\beta\tilde{\theta}_2)^{-1}$, and $\tilde{\eta}_{t+1} = \beta^{-1}\eta_{t+1} - (\beta\tilde{\theta}_2)^{-1}\epsilon_{1t}$. $\overline{\Delta} = (1 - \beta F)$, where F is the forward-shift operator.

In the steady state the model reduces to:

$$(8) \qquad a_1 + a_{11}L^s + a_{12}K^s + b_2W^s + b_3Q^s + b_{12}\exp(\theta_{12}Q^s) = 0 ,$$

where $a_1 = \overline{\alpha}_1 + \overline{\theta}_2\theta_{11}(1 - \beta)$; $a_{11} = \overline{\alpha}_{11} - \beta^{-1} - 1$; $a_{12} = \overline{\alpha}_{12}$; $b_2 = \overline{\theta}_2$; $b_3 = \overline{\theta}_3(1 - \beta)$; and $b_{12} = \overline{\theta}_2\theta_{12}(1 - \beta)$. Note that with the requirement that $a_{11} > 0$ the steady-state results make economic sense. That $b_2 > 0$ implies a downward-sloping long-run labor-demand curve. That $b_3 > 0$ implies the reasonable long-run result that, other things equal, additional quits generate higher user—costs of labor, thus raising the full cost of employment and reducing the number of workers demanded.

Equations (6)—(8) describe the solution and estimating equation for Model III. Model II is described by the same equations, but with $\theta_{12} = 0$. These equations also describe Model I if $\theta_{11} = 0$ as well.

III. Description of the Data

The estimation of these models is possible because until 1982 the United States collected monthly establishment—based data on flows of workers. Thus for each plant in the survey data were collected on accessions, divided into new hires, rehires and other accessions (mainly transfers between plants of the same firm, and returning military personnel), and separations, consisting of layoffs, quits and other separations (mainly workers discharged for cause).[4] We stress that these are gross flows of workers. They are entirely different data, and measure entirely different concepts, from data on changes in the level of employment resulting from flows of jobs as some firms expand and others contract (analyzed for a number of countries by, e.g. Wedervang, 1965; Davis and Haltiwanger, 1992).

Rehires are well described as a constant fraction of recent layoffs, and other separations appear to be a small constant fraction of new hires (Hamermesh, 1969). That being so, and ignoring the tiny flows of other accessions, we can estimate the models under the reasonable assumption that the identity in (1) is a good description of the link between net and gross changes in employment.

We estimate all the models for 1961:I–1981.IV, the last period for which the turnover data are available. The quit rate, and the number of workers implied by it, is the three–month sum of the monthly flows of quits. All the other variables are quarterly averages of monthly series. L is total manufacturing employment based on the monthly establishment survey. The forcing variable W is represented by real manufacturing compensation per hour paid for, an appropriately broader measure than hourly earnings. The capital stock is represented by a time trend.

In order to apply GMM to obtain parameter estimates of the first–order necessary conditions we collected a set of instrumental variables. To allow for the autocorrelation of the disturbances in the form of a first–order moving average the instruments must be lagged at least two periods. The instruments used are: Two– and three–period lagged changes in W and L; three–period lagged changes in Q, the Producer Price Index (PPI) and output; a three–period lagged term in manufacturing gross investment in structures and equipment; four–period lagged terms in Q, output, the civilian unemployment rate, the PPI, and the rate of capacity utilization; a constant and a time trend. Manufacturing output is measured by the Department of Commerce's series on manufacturing shipments minus the change in inventories of final goods. Capacity utilization is the FRB index for manufacturing, and gross investment is from the NIPA data. Seasonally adjusted versions of all of the variables are used in the estimation.[5] The discounting parameter β was set equal to 0.98 prior to estimation.

IV. Results

Table 1 presents GMM estimates of the adjustment cost parameters for Models I–III, as well as goodness-of-fit indicators such as the $\bar{R}^2$ and the standard deviation of the residuals (s.e.). Before discussing the economic meaning of the parameters, we first need to investigate the statistical properties of the models. A necessary condition for GMM to produce consistent parameter estimates is that $E_t\{\tilde{\eta}_{t+1}\} = 0$ and $\tilde{\eta}_{t+1}$ is stationary. The Sargan–Bhargava statistic (1983), which tests for the presence of a unit root in the residuals, indicates no significant non-stationarity in any of the residual series. The diagonal of Table 2 presents p-values of Hansen's J-statistic testing the over-identifying restrictions.[6] None of these restrictions is rejected at the 1-percent level for any of the estimated equations, although the J-test of Model I with $\theta_3 = 0$ does reject the orthogonality hypothesis at the 5-percent level.

The upper triangle of Table 2 presents Gallant's (1987) likelihood-ratio type statistic for testing nested restrictions. For example, to test whether quits matter in Model I we can look either at the asymptotic t-statistic on θ_3 of Model I in Table 1, or, alternatively, at the p-value on H_0: $\theta_3 = 0$ in the second row of Table 2. The estimate of θ_3 is positive and significant in Model

TABLE 1 : GMM Estimates of the Adjustment Cost Parameters *

	MODEL I		MODEL II		MODEL III
	(θ_2)	(θ_2, θ_3)	(θ_2, θ_{11})	$(\theta_2, \theta_3, \theta_{11})$	$(\theta_2, \theta_3, \theta_{11}, \theta_{12})$
θ_2	33.526 (2.171)	11.968 (2.159)	54.545 (2.206)	51.731 (2.084)	51.996 (2.214)
θ_3	0	0.430 (2.256)	0	-0.100 (-0.677)	1.297 (1.651)
θ_{11}			-7.015 (-4.682)	-6.869 (-4.462)	-6.959 (-4.725)
θ_{12}					-1.198 (-3.667)

* Asymptotic t-values in parentheses.

Goodness-of-Fit Indicators

	MODEL I		MODEL II		MODEL III
$\bar{R}^2$	0.97509	0.97596	0.99947	0.99946	0.99955
s.e.	0.01102	0.01003	0.00159	0.00161	0.00147
SB	2.214	2.076	2.690	2.647	2.425

I, not significantly different from zero in model II, but positive and significant at the 10-percent level in Model III.[7] The estimates of θ_2 are significantly positive in all five equations. Positive estimates of θ_2 and θ_3 are sufficient -- and in the LQ-model necessary -- conditions for strictly convex adjustment costs and point at the usual quasi-fixity of labor. We indeed find that both the costs of changing the net level of employment and costs of replacement are important for U.S. manufacturing, especially if one estimates the usual linear-quadratic model.

The estimates of θ_{11} appearing in Models II and III are significantly negative. The intuition is that negative net changes in employment are more costly than positive net changes. This result is consistent with earlier findings for manufacturing nonproduction workers in the Netherlands and the U.K., but not with findings for production workers (Pfann and Palm, 1992). Moreover, within the context of the representative firm, the observed asymmetry of the responses to $\Delta L_t + Q_t$ ($\theta_{12} < 0$) implies that it is more costly to lay off employees than it is to hire new workers.

The standard asymmetry apparent in the data is that reductions in employment occur faster than increases (implying that the asymmetry parameters are positive). The cyclicality of quits reduces this kind of asymmetry by slowing cuts in employment. Here we generate estimates of structural parameters, which need not yield the same implications as reduced-form coefficients that depend both on the structural parameters and on the

TABLE 2 : Tests of Parameter and Overidentifying Restrictions *

MODEL I		MODEL II		MODEL III
(θ_2)	(θ_2, θ_3)	(θ_2, θ_{11})	$(\theta_2, \theta_3, \theta_{11})$	$(\theta_2, \theta_3, \theta_{11}, \theta_{12})$
p = .027 J1 11	p = .024 $H_0 : \theta_3 = 0$ 1	p = .000 $H_0 : \theta_{11} = 0$ 1	p = .000 $H_0 : \theta_3 = \theta_{11} = 0$ 2	p = .000 $H_0 : \theta_3 = \theta_{11} = \theta_{12} = 0$ 3
	p = .085 J2 10		p = .000 $H_0 : \theta_{11} = 0$ 1	p = .000 $H_0 : \theta_{11} = \theta_{12} = 0$ 2
		p = .143 J3 10	p = .513 $H_0 : \theta_3 = 0$ 1	p = .211 $H_0 : \theta_3 = \theta_{12} = 0$ 2
			p = .127 J4 9	p = .107 $H_0 : \theta_{12} = 0$ 1
				p = .081 J5 8

* Each box lists the significance level, the hypothesis and the degrees of freedom.

temporal pattern of the random shocks. In any case, since our estimates are based on data covering all employees, our results imply that employers' concerns about forfeiting shared investments in firm-specific human capital dominate their propensity to avoid incurring hiring costs. This is consistent with the arguments of Bentolila and Bertola (1990).

V. Conclusions, and Implications for Dynamic Labor Demand

We have developed and estimated a model of labor demand that accounts for dynamics arising from the costs of both net and gross changes in employment. The estimates suggest that observed lags in the demand for labor at the aggregate level arise from both sources of costs. Moreover, both produce significant asymmetric adjustment.

Though we have constructed and estimated a model that rests on data that were only available for one economy and that are no longer being collected, our findings are not merely an historical curiosum. We have shown that a measure of voluntary turnover does belong in the estimation of a formal model of dynamic labor demand, and that our understanding of these dynamics is diminished by the absence of these data. Most important, the estimates of Model I with and without the constraint $\theta_3 = 0$ show that much of the slow adjustment that is attributed to the costs of changing the level of employment in fact results from the costs of replacing workers who have quit. This means that the interpretation of lag parameters, including comparisons of employment lags across economies, must be done with great care.[8]

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FOOTNOTES

1. Whether quits are truly exogenous to the employer is an extremely complex issue. That they are logically separate categories that can usefully be distinguished, though, is clear (McLaughlin, 1991). By using the published data we implicitly rely on each employer's classification of separated workers into the two main categories, quits and layoffs.

2. A simple regression of log employment on the quit rate and a time trend yields:

$$\log(L) = \underset{(307.6)}{1.17} + \underset{(25.45)}{.924}{*}T(x1000) + \underset{(18.83)}{1.20}{*}Q, \quad \bar{R}^2 = .931,$$

where Q is the seasonally adjusted quarterly quit rate, and we list t−statistics here and throughout the paper in parentheses. The simple correlation of log (L) and Q is .63. For small industries from 1958−1966 the elasticity of the quit rate with respect to aggregate unemployment was −2.6 (Hamermesh, 1969).

3. The two studies only have access to time series on accessions and separations, the latter of which include voluntary quits and the endogenous layoffs. They are thus incapable, as Burgess and Nickell explicitly recognize, of linking their modeling of the representative firm's choices to the estimates.

4. See, for example, the description in the statistical section of any issue of the Monthly Labor Review before 1983.

5. Standardized forms of each variable, $Z^* = (Z/\bar{Z}) - 1$, where $\bar{Z}$ is the mean of Z, were used in the estimation.

6. This is a χ^2−test with degrees of freedom equal to the number of instruments minus the number of estimated parameters.

7. In Model III the other structural parameters are:

$$\alpha_1 = \underset{(1.57)}{8.825};\ \alpha_{11} = \underset{(.79)}{.208};\ \alpha_{12} = \underset{(4.90)}{.00307}.$$

While we cannot be sure that A is positive definite, since we do not estimate α_{22}, the positive point estimates of the other three parameters are good indications that this fundamental description of technology is valid in these data.

8. A recent example of such comparisons is Abraham and Houseman (1989).