COWLES FOUNDATION FOR RESEARCH IN ECONOMICS

AT YALE UNIVERSITY

Box 2125, Yale Station
New Haven, Connecticut 06520

COWLES FOUNDATION DISCUSSION PAPER NO. 848

AN AGGREGATIVE DISEQUILIBRIUM MODEL OF THE U.S. LABOUR MARKET

by

Vassilis Argyrou Hajivassiliou

July 1987

An Aggregative Disequilibrium Model of the U.S. Labour Market

by

Vassilis Argyrou Hajivassiliou*

Cowles Foundation for Research in Economics

Yale University

July 1985

Revised June 1987

ABSTRACT

A model is presented in which aggregation over microsectors, each in differing extent of disequilibrium, has implications analogous to the standard single aggregate sector switching disequilibrium model. Empirical implementation of the model of this paper is less involved than estimation of the standard model. Hence the approach here may be seen both as providing an underlying micro justification for the switching disequilibrium model, and as a computationally simpler (though statistically less efficient) technique. The model is estimated from post-war labour market quarterly data for the U.S. manufacturing sector. We find the supply side more satisfactorily determined than in past disequilibrium studies.

---

*I would like to thank Olivier Blanchard, Stanley Fischer, Daniel McFadden, Whitney Newey, Richard Quandt, Robert Solow, John Sutton, and James Tobin for very helpful comments and suggestions. The author retains full responsibility for any errors and shortcomings.

An Aggregative Disequilibrium Model of the U.S. Labour Market

1. Introduction

The equilibrium hypothesis plays a crucial role in economics. A leading example is the rational expectations school associated with the names of Lucas, Sargent and Wallace. Though the alternative of "disequilibrium" modelling literature has had a long tradition in theoretical economics, attempts at disequilibrium econometrics have taken off only from the early seventies with the pioneering works of Fair and Jaffee (1972) and Maddala and Nelson (1974).

The leading disequilibrium modelling technique builds on the postulate of rationing in the face of voluntary exchange and so yields the short-side or minimum rule (see, for example, Barro and Grossman (1971)). The switching regression model has been extensively developed by Maddala and Nelson (1974), Goldfeld and Quandt (1975, 1978, 1979), Quandt (1982), and others. The basic model is:

$$D_t = X_t^d \beta_t^d + \varepsilon_t \tag{1}$$

$$S_t = X_t^s \beta_t^s + u_t \tag{2}$$

$$Q_t = \min(D_t, S_t) \tag{3}$$

where Q is transacted quantity. More recent extensions involve dynamic models and serially correlated errors with observations on regimes (Laffont and Monfort (1979)) or without regime classification information (Quandt (1981)), Lee (1984a)). It must be noted that there are significant problems with the switching model:

1. Especially when applied directly on aggregate data as in the case of Rosen and Quandt (1978), there is the very valid objection (e.g., Malinvaud

(1980)) as to the plausibility of an economy-wide sector, like the labour market, or of the whole economy itself switching discretely between excess demand and excess supply regimes.

2. In the face of endogenous switching, the method of maximum likelihood may be hindered by severe problems involved with the parameter space containing singularities or the likelihood function being unbounded (see Quandt (1981)).

3. The switching model imposes sometimes astonishing computational burdens. See, for example, Quandt (1982) for treating the mere presence of first order autoregressive processes in the equation errors. Though this is less of a concern with the recent surge in computer technology, it does explain partly why disequilibrium econometric studies remain so sparse.

Attempts at overcoming criticism 1, include the "fuzzy switching" model where an error is appended to the minimum condition (see Bowden (1978)), Quandt (1982)), and aggregation models such as the ones independently proposed by Muellbauer (1978) and Malinvaud (1980). The problem of aggregating micro markets in disequilibrium has also attracted the attention of Kooiman (1984), Lambert (1984) and Monfort (1982). My paper will pursue the aggregation approach. The resulting estimation technique is considerably better behaved computationally, especially in handling serial correlation in the unobservables.

In section 2, we present the aggregation model and draw out its implications about aggregate employment. Section 3 works out the main estimable version of the model under specific distributional assumptions. Various data and econometric issues are discussed in section 4. The specific modelling of notional demand and supply functions is described in section 5. Section 6 presents the empirical results of this study, which are particularly more successful for the supply side, compared to past disequilibrium studies.

2. Theory

Think of the economy or the relevant market as consisting of J sectors assumed of equal size for simplicity; otherwise simple scale factors are needed. Within each sector, model (1)-(3) applies. $X_t^d\beta_t^d$ and $X_t^s\beta_t^s$ are to be thought of as mean (economy wide) demands and supplies (D_t^*, S_t^*), and

$$\begin{aligned}
&(1)' \qquad D_t^j = D_t^* + \varepsilon_t^j \\
&(2)' \qquad S_t^j = S_t^* + u_t^j \qquad\qquad j=1,\ldots,J \\
&(3)' \qquad Q_t^j = \min\,(D_t^j, S_t^j)
\end{aligned}$$

where ε^j, u^j are sector-specific demand and supply shocks. Dropping the time subscript for simplicity, we aggregate over sectors on the basis of a postulated distribution function of the shocks $F(\varepsilon,u)$, conditioning on any temporal (economy-wide) randomness included in D_t^* and S_t^*. (Note the lack of a j index, implying that the sector-specific shocks are drawn from an identical distribution, the drawings assumed independent across j).[1] Given condition (3)', transacted quantity in the aggregate is given by

$$(4) \qquad Q = \sum_{\text{for } j\in\{D^j<S^j\}} D^j \; + \sum_{\text{for } j\in\{D^j>S^j\}} S^j$$

Let A be the set of sectors that are in excess demand, i.e., $\{j\,|D^j>S^j\} \equiv A$ and let B be its complement $\{j\,|D^j<S^j\} \equiv B$. Making the assumption of many markets so as to give an approximately continuous distribution $F(\cdot,\cdot)$, we can replace summations by integrals to obtain

$$\begin{aligned}
Q &= \iint_B D^j dF(\varepsilon,u) + \iint_A S^j dF(\varepsilon,u) \\
&= \iint_B (D^* + \varepsilon^j) dF(\varepsilon,u) + \iint_A (S^* + u^j) dF(\varepsilon,u) \\
&= \iint_B D^* dF(\varepsilon,u) + \iint_A S^* dF(\varepsilon,u) + \iint_B \varepsilon\, dF(\varepsilon,u) + \iint_A u\, dF(\varepsilon,u) \\
&= D^* \iint_B dF(\varepsilon,u) + S^* \iint_A dF(\varepsilon,u) + \varepsilon^* + u^*
\end{aligned}$$

$$(4') \qquad Q = D^*(1-\pi) + S^*\pi + \varepsilon^* + u^*,$$

where π is the proportion of the sectors that are in excess demand. ε^* denotes the average demand shock in the sectors that are in excess supply and u^* the average supply shock in sectors with excess demand. π is clearly dependent upon the magnitude of total excess demand (D^*-S^*) through the density function $dF(\varepsilon,u)$, and so are ε^* and u^*.[2] Further note that ε^* and u^* are sufficiently negative so that Q is always on the inside of the min(D,S) wedge (see Figure 1). Intuitively this is because, since all the micro markets are rationed, you will always be staying <u>shorter</u> than the aggregate short side rule, except in the limiting case of <u>all</u> markets lying in the <u>same</u> regime. This further "loss" will clearly be at a maximum when there is the greatest degree of "imbalance" in the economy indicated by a value of $\pi = 1/2$.

Such an idea of a "stochastic nature of macro equilibrium," whereby the individual micro sectors may be in very different extents of excess demand, is used by Tobin (1972) and formalized by Sutton (1981) to generate a model of inflation. It seems appealing economically to postulate that some sectors of the economy (or of the market under analysis) may be in excess demand (computer industry? computer specialists?) while others in excess supply (auto industry and workers?) with the <u>proportion</u> of the two categories of sectors depending on the extent of the whole economy or market-wide strength of demand versus supply. This is also very well substantiated by survey studies (see Sutton (1981) and Bouissou <u>et</u> <u>al.</u> (1986) with respect to data on firms). Muellbauer (1978) attempts to augment the model further by allowing for spillovers among sectors as time goes by. Here I will proceed with equation (4') and discuss estimable versions of it by making standard assumptions about $F(\varepsilon,u)$.

We first note the following equivalence between equation (4') and the standard switching regression model formulated in expected value terms. Consider equations (1)-(3) for the whole market and denote by $\bar{D}_t$ and $\bar{S}_t$ the deterministic parts of D_t and S_t. Let A denote the event $\{D_t > S_t\}$ with associated probability $P_A \equiv P$, and B its complement. (i.e., $A \equiv \{D_t > S_t\} \equiv \{\bar{D} - \bar{S} > u - \varepsilon\}$). Taking expectations of equation (3):

$$\begin{aligned} EQ &= E\min(D,S) = E(D|B)P(B) + E(S|A)\cdot P(A) \\ &= E(\bar{D} + \varepsilon|B)\cdot(1-P) + E(\bar{S} + u|A)\cdot P \qquad (5) \\ EQ &= \bar{D}\cdot(1-P) + \bar{S}\cdot P + E(\varepsilon|B)\cdot(1-P) + E(u|A)\cdot P \end{aligned}$$

This formulation makes it apparent that a direct analogy exists between the switching model (equation (5)) and the aggregation approach (equation (4')). First, note that (4') is still conditioned on the temporal (economy-wide) component of randomness and so it is Q and not EQ that is the dependent variable of the model. We will have precisely the same RHS expression and EQ as the dependent variable in the section 3 below, once we assume further that taking expectations over the temporal component of randomness from equation (4'), factors out appropriately. A sufficient condition for this is that $D_t^* = \bar{D}_t + \theta_t$, $S_t^* = \bar{S}_t + \theta_t$, θ_t being the economy-wide temporal random component, which is assumed common to both demand and supply sides.[3]

Hence we see that under appropriate conditions, aggregating over micro-sectors in disequilibrium predicts that transacted quantity will be given by the total mean demand and demand shocks weighted by the <u>proportion</u> of sectors in excess supply, <u>plus</u> the total mean supply and supply shocks weighted by the proportion of sectors in excess demand. In an exactly analogous manner, in the case of the switching approach on a single market, expected transacted quantity is given by the convex combination of expected demand and expected demand-error, given excess supply on one hand, and the expected supply plus the expected supply-error, given excess demand on the other. The weighting factor is in this case the <u>probability</u> of the single sector being in excess demand.

Further, it should be noted that equation (5) can, as a general principle, be written in actual Q terms as

(5') $\quad Q = EQ + v$,

where v is a heteroskedastic error term.[4] An important point is that equation (5') is computationally more tractable to estimate than the switching model (1) to (3), as it only requires non-linear least squares estimation (NLLS), readily available in most econometric packages.

To recapitulate, the aggregation approach presented in this paper yields two main results: First, the model (5)-(5') is obtained which may be thought of either as a computationally less complicated approach than switching ML at the cost of some statistical efficiency, or as an economically meaningful model that hypothesizes aggregation over various micro sectors in varying states of disequilibrium. The latter view provides an underlying story for the switching model applied on a single sector, with the appropriate reinterpretations. This equivalence should be borne in mind throughout the paper while discussion will proceed in terms of switching model formulated in terms of expected values.[5]

3. The Model in the Case of Normality of the Errors

Assume, as is customary in the literature, that ε and u are jointly normal with covariance matrix $\begin{bmatrix} \sigma_\varepsilon^2 & \sigma_{\varepsilon u} \\ \sigma_{\varepsilon u} & \sigma_u^2 \end{bmatrix}$ and i.i.d. across time. Simple manipulation of model (1)-(3) yields:

(1) $\quad D = X^d\beta^d + \varepsilon$

(6a) $\quad S-D = (X^s\beta^s - X^d\beta^d) + u - \varepsilon \equiv (S^*-D^*) + \eta$

(6b) $\quad D-S = (X^d\beta^d - X^s\beta^s) + \varepsilon - u \equiv - (S^*-D^*) + w, \; w \equiv - \eta.$

By the minimum condition (3), Q = D when S-D > 0 and Q = S when D-S > 0. Hence applying results from Heckman (1979) we obtain the expressions:

$$(7)\qquad E(D|S\text{-}D > 0) = X^d\beta^d + \frac{\sigma_{\varepsilon\eta}}{\sqrt{\sigma}_{\eta\eta}} \frac{\phi(\frac{D^*-S^*}{\sqrt{\sigma}_{\eta\eta}})}{1-\Phi(\frac{D^*-S^*}{\sqrt{\sigma}_{\eta\eta}})}$$

and

$$(8)\qquad E(S|D\text{-}S > 0) = X^s\beta^s + \frac{\sigma_{uw}}{\sqrt{\sigma}_{ww}} \frac{\phi\ (\frac{S^*-D^*}{\sqrt{\sigma}_{ww}})}{1-\Phi(\frac{S^*-D^*}{\sqrt{\sigma}_{ww}})},$$

where $\phi(\cdot)$ and $\Phi(\cdot)$ denote the standard normal density and cumulative distribution functions respectively. These results are derived explicitly in Appendix A by using standard properties of the multivariate normal distribution, as given in Johnson and Kotz (1972). Noting that $\eta \equiv - w \equiv u - \varepsilon$ and that $F(x) = 1-F(-x)$, we get

$$(7')\qquad E(D|A^c) = X^d\beta^d - \frac{(\sigma_\varepsilon^2-\sigma_{\varepsilon u})}{\sigma_\eta} \frac{\phi\ (\frac{D^*-S^*}{\sigma_\eta})}{1-\Phi(\frac{D^*-S^*}{\sigma_\eta})}$$

and

$$(8')\qquad E(S|A) = X^s\beta^s - \frac{(\sigma_u^2-\sigma_{\varepsilon u})}{\sigma_\eta} \frac{\phi\ (\frac{D^*-S^*}{\sigma_\eta})}{\Phi(\frac{D^*-S^*}{\sigma_\eta})},$$

where $\sigma_\eta^2 = \sigma_\varepsilon^2 + \sigma_u^2 - 2\sigma_{\varepsilon u}$ and, in our previous terminology, P(A) = P(Excess Demand) $= \Phi(\frac{D^*-S^*}{\sigma_\eta})$. Therefore under normality, equation (5) reads:

$$(9)\qquad EQ = D^*\cdot[1-\Phi(\frac{D^*-S^*}{\sigma_\eta})] + S^*\cdot\Phi(\frac{D^*-S^*}{\sigma_\eta}) - \sigma_\eta\ \phi(\frac{D^*-S^*}{\sigma_\eta}) .$$

This is the equation to be estimated by NLLS with a White consistent covariance correction.[6,7] Note that equation (9) is obtained because no information on regimes is assumed - the conditional expectations of Q on regime 1 or 2 being observed are weighted by the appropriate probability so as to get the unconditionally expected Q.

4. Data and Preliminary Issues

The market under examination is the aggregate labour market in the U.S. manufacturing sector and the period covered is 1948QI-1983QII. The data used are <u>seasonally</u> <u>unadjusted</u> quarterly series contained in the CITIBANK data base. Details are provided in Appendix C. As a point of reference, I chose a disequilibrium study of the aggregate U.S. labour market by Rosen and Quandt (1978) and its subsequent corrections and revisions (Romer (1981), Quandt (1981)), which use annual data from 1930 to 1973. The basic model specifies log linear demand and supply, plus the short-side (minimum) rationing rule.

Before discussing the particular specifications of (log)D_t and (log)S_t tried, the following two general issues are noted:

1. <u>Endogeneity of the Wage?</u>

The initial Rosen and Quandt (1978) study modelled the wage in an explicitly endogenous way by appending a wage-adjustment equation to the three basic equations of the switching model:

$$(10) \qquad \ell nW_t - \ell nW_{t-1} = \gamma_1(\ell nL_t^d - \ell nL_t^s) + \gamma_2 V_t + \xi_t$$

where for V_t, 3 alternatives were tried: (i) a vector of ones, (ii) the unionization rate and (iii) the change in the unionization rate. MLE was carried out on equations (1), (2), (3) and (10). In his latter study, Quandt (1981) decided to drop equation (10) essentially for computational reasons. I chose Quandt's latter approach with a further consideration in mind:

Treating the wage at least preliminarily as predetermined may be justified as a first approximation in a (quarterly) labour market context and in view of evidence on Phillips (wages) equations where contemporaneous excess demand for labour terms are found to have very insignificant effect on wage adjustment (see Godley and Nordhaus (1972) for the U.K., Nordhaus (1976) for the U.S.). Some justification may also be gleaned from evidence that employment does not Granger-cause wages (Neftci (1978)) and also wages do not Granger-cause employment (Geary and Kennan (1982)). Though Granger non causality does not bear any one-to-one relation to econometric exogeneity and/or predeterminedness (on this see Hausman (1983) and Engle, Hendry and Richard (1983)), it seems that treating the wage as predetermined as part of a disequilibrium hypothesis may not be too unreasonable. Therefore in what follows, I implicitly postulate instead of (10), an equation of the form

$$(10') \qquad \ln W_t - \ln W_{t-1} = \sum_{i=1}^{k} \gamma_{1i} \, (\ln L^d_{t-i} - \ln L^s_{t-i}) + V_t \gamma_2 + \xi_t,$$

where ξ is serially uncorrelated and contemporaneously uncorrelated from ε and u, with V a vector of relevant exogenous or predetermined variables. Exogeneity tests in maximum likelihood disequilibrium models, developed in Hajivassiliou (1986b), will be applied in the estimation section below. What is of crucial importance for our purpose is that no contemporaneous excess D term appears, a reasonable assumption with quarterly data[8,9].

2. Treatment of Serial Correlation

It is a well known fact of economic life that aggregate data exhibit very strong serial correlation. Hence it is somewhat surprising that Quandt (1981), allowing each error to be an AR(1), finds the first order autocorrelation coefficients both to be individually statistically insignificant at any plausible level of significance and certainly of inconsequential estimated magnitudes. Given that our data are quarterly, however, serial correlation is more likely. Robinson (1982) suggests and Hajivassiliou (1986a) proves that serial correlation in the disequilibrium switching model with normally distributed errors does not pose consistency problems for MLE, so the hope is that it will not be too critical, if serial correlation is not overcome fully.[10] Unfortunately, correct treatment of serial correlation in non linear models in general and in disequilibrium models in particular is extremely intractable (see inter alia Quandt (1981) and Lee (1984a)). Possible approaches are discussed in Hajivassiliou (1985). Serial correlation can be introduced in the aggregative disequilibrium model, (5') and (9), in the following manageable way: allowing the error v_t, which corresponds to the economy-wide temporal random component, to be autocorrelated, the required estimation procedure is NLLS with an autoregressive error. This possibility is examined below.

A related issue that we address carefully here is the presence of seasonality. Given the non-linearities of the model, one should use seasonally unadjusted data and carry out deseasonalization by the inclusion of dummy variables on both D and S sides, simultaneously with the estimation of the structural parameters of the model. Even though this is a computationally very demanding task, we follow this approach here. To our knowledge, the problems with using deseasonalized data are not acknowledged in most non-linear studies (see, for example, Hansen and Singleton (1982)).[11] This decision, to work with seasonally unadjusted series and carry out the deseasonalization simultaneously with the estimation, unfortunately necessitated restricting attention to the manufacturing sector.

5. Specific Modelling of L^D, L^S

1. Demand: The most general specification tried was the following (in logarithmic form):

$$(13)\quad L_t^D = a_0 + a_1'\{\text{seasonal dummies}\} + a_2 \cdot T + a_3 \cdot W_t + a_4 \cdot Y_t + a_5 \cdot W_{t-1} + a_6 \cdot Y_{t-1} + a_7 \cdot W_t^* + \epsilon_t$$

This corresponds to the marginal productivity condition of cost minimization of the neoclassical firm. W is the (gross) real wage. The deflating was tried both by the CPI index and by the theoretically more relevant WPI deflator.[12]

Y is industrial production and T is a time trend supposed to capture secular changes in the capital stock and other long term trends in technology. In a disequilibrium model, Y is relevant as a quantity signal for a firm which is possibly rationed on the goods market, as it also is in an imperfect competition setting since it affects the marginal revenue product of labour. Y_t is treated as predetermined -- through exogenous government policy and autonomous aggregate demand shocks.[13]

W_{t-1} and Y_{t-1} were tried both for remedying some effects of serial correlation in the errors, and as possible persistence effects due to adjustment costs in employment decisions. W_t^* is a measure of the expected (at t) wage to rule next period. (For a model which shows that the rational firm facing adjustment costs with respect to employment decisions will be both backward (W_{t-1}) and forward (W_t^*) looking, see Sargent (1978).) W_t^* was somewhat mechanically constructed by the ARMA Box-Jenkins (1970) methodology using the TSP V4.0 computer package. See Table 1. The most satisfactory parsimonious ARMA specification appears to be an AR(2) with drift, a time trend and seasonals. [Evidence on this issue is given also in Altonji (1982) and in Altonji and Ashenfelter (1980)]. W_t^* are the predictions from this fitted equation. There are two issues with this construction:

(i) Using the fitted values obtained from an equation estimated on the whole sample implicitly assumes that people forecast on the basis of information unavailable at the time. Hence it is theoretically more justifiable to use the

Kalman updating technique (see, e.g., Harvey (1981)). However, Barro (1977, 1978) reports that it mattered very little for all practical purposes whether the above approach was adopted, or whether a separate updated regression was run using only observations predating the date of the forecast.

(ii) The rational expectations hypothesis argues explicitly against the postulate that people form expectations solely on the basis of the past history of the variables involved. But in view of the evidence on how well fairly simple ARMA models forecast, sometimes even better than large structural equation models [on this see Granger and Newbold (1977), Chapter 8], taking W^* as predictions from a simple ARMA seems to be a plausible working hypothesis.

2. <u>Supply</u>: Here the most general specification (again in logarithmic form) was:

$$
\begin{aligned}
(14)\quad L_t^s = {} & \beta_0 + \beta_1'\{\text{seasonal dummies}\} + \beta_2 \cdot X_t + \beta_3 \cdot X_{t-1} + \beta_4 \cdot R_t + \beta_5 \cdot POTL_t \\
& + \beta_6 \cdot POTL_{t-1} + \beta_7 \cdot X_t^* + \beta_8 \cdot RELW_t + \beta_9 \cdot RELW_{t-1} + u_t
\end{aligned}
$$

Such an equation may be obtained from the intertemporal maximization problem of a representative worker. X_t is the <u>after-tax</u> (net) wage he/she receives deflated by the consumer price index. X_t^* is the worker's expectation of the future net wage, constructed once more using the most parsimonious ARMA specification for X_t , in this case a random walk plus constant, a trend and seasonals. See Table 1. X_{t-1} may be rigorously justified by habit-formation which makes the utility function intertemporally nonseparable.

R_t is the <u>ex ante real interest rate</u> which plays an important role in the intertemporal labour substitution hypothesis, in that it affects the choice at the margin of working more now and saving versus taking more leisure and borrowing or running down assets. The nominal interest rate variable used was the 3-month Treasury bill rate; this may be problematic, since Treasury bills were not readily accessible to the average worker over the study period. However, most studies on the intertemporal consumer-worker optimization use this and/or similar variables (see e.g., Hansen and Singleton (1982), Mankiw, Rotemberg, and Summers (1985)) and the results appear to be quite robust as to the choice of nominal R-variable.

Note that one may be facing the possible inconsistency of the workers being rationed in the labour market (which is the crux of the idea of disequilibrium in the labour market), while having perfect access to the capital market at the competitive price R. It is more plausible to expect quantity constraints in the labour market to go together with credit-rationing in the capital market.

One should also acknowledge that the forecasting equation for the inflation rate (see Table 1) may be very different from the forecasting mechanisms people were actually using. But this is a general problem in the treatment of expectations (which clearly applies also to the W^* and X^* constructions already encountered).

$POTL_t$ is a measure of the full potential man-hours that could be worked if all population of working age were attracted in the manufacturing labour force. Rosen and Quandt (1978) and the subsequent studies mostly use as $POTL_t$ the total civilian population of working age <u>times</u> the average man-hours worked in the economy, hence making the implicit assumption that if more people were attracted in the labour force, they would be working the same hours as the hitherto employed. This measure of POTL, however, creates a serious econometric problem: if average hours worked are constructed, as to be expected, by dividing the actual man-hours worked by the civilian labour force, we have directly introduced <u>endogeneity</u> on the RHS of the equation to be estimated for labour employment, in that a (non-linear) function of the dependent variable has been entered directly on the RHS. To avoid this problem, POTL is measured here by the total civilian labour force. $RELW_t$ is entered as a measure of the wage of production workers in manufacturing, relative to the one of such workers in the overall labour market. This attempts to capture the relative attractiveness of seeking employment in the manufacturing sector, as opposed to the "average" market.[14]

6. Empirical Results

Table 2 presents the first estimated versions of equation (9). See Appendix C for the exact definitions of variables. AGD refers to the aggregated disequilibrium model (eq. (9)) estimated by NLLS, with heteroskedasticity consistent covariance matrix. The signs of the coefficients come out as predicted by theory and as also reported by similar studies (see, e.g., Rosen and Quandt (1978), Romer (1981) and Quandt (1981)). The wage effect is unambiguously negative on the demand. Magnitudes of the coefficients cannot be compared to the ones of the aforementioned studies as they use different data (from different sources and of different periodicity). However, the positive effect of real industrial production on labour demand is strongly confirmed (asymptotic t statistics greater than 5). In general, a positive elasticity of supply with respect to the (after tax) wage is found. In some estimations, the contemporaneous after tax wage comes out negative, but statistically insignificant. While this agrees with the studies cited above, and while economic theory leaves this elasticity unsigned, it is interesting to note that the generally positive sign of the one-period lag of this variable suggests the existence of adjustment lags. Changes in the potential labour force have a positive contemporaneous effect on the labour supply (though not of such pervasive significance: t rarely exceeds 2). The lagged effects of the potential labour force are not well determined in general. As predicted by theory, the relative wage in the manufacturing sector compared to the overall economy one, affects strongly in a positive way the supply of labour to the manufacturing sector. The SIGH parameter which is the standard deviation of excess demand standardizing the probabilities expressions (σ_{η} in equation (9)) comes out positive as it should.

Trend variables supposed to capture capital stock movements and long-term productivity-technology changes in labour demand, and changes in participation rates in labour supply, are also tried. These have coefficients generally of the wrong sign (positive on demand) and rarely very robust to specification changes. The same applies to quarterly dummy variables that model for seasonality. This instability may simply point out the already noted computational problems, of dealing with seasonality in such non-linear models. Seasonality was found statistically strongly significant, with an LR statistic in AGD1 of 47.32, against a $\chi^2_{1\%}(6) = 16.8$.

Estimates were also obtained by the MLE switching-regimes disequilibrium model (Fair and Jaffee (1972)) and are denoted by SNOR (for Switching with No-Observations-on-Regimes). The closeness of the MLE results to AGD comparable versions obtained from our model eq(9), the very similar performance and fit of these two approaches to be discussed below, and the Monte-Carlo results in Quandt (1986), provide the basic support for the aggregation approach of this paper as a viable alternative to and/or an underlying model for the standard ML disequilibrium approach. Versions 2 of each approach suppress the real interest rate on supply. The sign of the real interest rate effect comes out positive as the intertemporal labour substitution hypothesis predicts, and is in general statistically very highly significant. To test the "no fiscal illusion" hypothesis, the (log of the) real wage and the (log of the) (1-TH) variable were entered separately with different coefficients. Under the "no fiscal illusion" hypothesis, the equality of these two coefficients should not be rejected, which indeed is the case (with a χ^2 of 1.231).

We further examined the significance of the appropriate measures of expected wage on the labour demand and supply. We were unable to carry out such tests because of very strong collinearity between current and lagged wages, and the expected measures. These collinearities seem unlikely to have been reduced by refinements of the forecasting equations. The estimates in Table 3 labelled AGD1SC were obtained by allowing the error v_t in eq. (9) to follow a (stationary) autoregressive process of order 1. The very high value of ρ (.87) and its strong statistical significance cast some doubt on the reported estimated standard errors, while not on the consistency of the parameter estimates.

We finally performed some specification tests on equation (9) of the AGD models. Since the standard deviation parameter σ appears both in the cumulative normal terms (probability of excess D in the major functional form of equation (9)), as well as in the conditional expectations (normal density) term, a potentially powerful test of specification is to let the two σ's differ, say σ_1 inside the c.d.f. and σ_2 inside the correction term, and test whether they differ significantly . Both tests did not reject the specification in this sense (the largest "LR" comes out at 1.366 for AGD1 against a $\chi^2(1)$ 10% critical = 2.71).[15] Furthermore, Table 3 reports some wage-exogeneity tests that employ the methodology in Hajivassiliou (1986b), applicable to MLE estimation of the SNOR model. To carry out such tests, one is required to obtain OLS estimates of the reduced-form (RF) equations of the variables suspected for endogeneity, and test the significance of these RF residuals when entered as additional variables on the D and S sides. In view of the high degree of serial correlation reported above, however, one should bear in mind that these tests are based on probably unreliable estimates of the covariance matrices. With this caveat in mind, we find that the SNOR1RW results reject marginally the exogeneity of real wage with a LR statistic of 14.8 against a critical value of $\chi^2(4) = 13.3$ at the 1% level of significance. SNOR1NW on the other hand does not reject the exogeneity of the nominal wage with a LR = 0.18, vs. a critical value of $\chi^2_{10\%}(2) = 4.6$. We hence view the evidence on this matter as mixed.

So as to further evaluate the estimation results we present two measures of fit between actual data and predictions of the model. First, employment predictions, using the "preferred" version AGD1, track employment very well. See the very high correlation figures between predicted and actual employment appearing in Tables 2 and 3. The model is able to reproduce quite accurately even <u>turning</u> points in the LE series -- the simple correlation between the actual LE series and the one predicted by estimated eq(9) is over 0.9. Figure 2 presents actual LE together with LE predicted by the AGD1 and SNOR1 models. Their good tracking ability should be obvious. As a second test of degree of fit, the estimated models were asked to predict the probability of excess demand at each time period (or the <u>proportion</u> of sectors in excess demand in our preferred interpretation) and then these predicted probabilities were examined for conformity to some well defined measure of "tightness" in the economy. To construct such a measure we first defined

XDY ≡ (real GNP - potential real GNP)/potential real GNP,

where potential real GNP is a series obtained from the Council of Economic Advisors (CEA). A second measure of slackness we tried was the level of unemployment (U).[16] Table 4 presents such correlations. PHAT refers to probabilities predicted by AGD1 while PROBH and PROBHC to ones predicted by the MLE model SNOR1. PROBH is given by $\Phi((D^*-S^*)/\sigma_\eta)$ evaluated at the estimated values, whereas PROBHC is the predicted probability <u>conditional</u> on the observed level of employment (i.e., Prob $(D_t - S_t > 0 | Q_t)$), which was shown by Lee (1984b) to minimize the total probability of misclassifying the two regimes.

All the predicted probabilities correlate very highly among themselves, and quite satisfactorily with both the XDY measure and the unemployment rate (-U). The .922 correlation between PROBH and PHAT provides further evidence for the closeness of the approach of this paper to the standard switching MLE model. In view of the strong evidence of a very close to unity root in the v_t errors, we present (annualized) values of PROBH predicted from the model formulated in <u>first differences</u>, and <u>changes</u> in XDY01 (the latter measure is simply changes in XDY transformed to lie between 0 and 1). The ability of the predicted probabilities series to track <u>turning</u> points in the <u>changes</u> in the tightness-of-the-economy measure is quite impressive. See Figure 3.

A further interesting feature of these ex-post "predictions" is that as a <u>rule</u>, the model estimated the excess supply "regime" as more likely overall, with predicted excess demand probabilities or proportions lying mostly <u>below</u> 0.5. In other words, even though the <u>changes</u> in tightness are followed impressively by the predictions, the model seems to tell us that the labour market is in <u>most</u> time periods in aggregate "slackness" rather than in "tightness". While this finding may simply indicate that our modelling of the demand side is better than the one for the supply, the evidence seems to confirm the Keynesian view of the world, where the economy fluctuates below a "ceiling" of full employment, rather than <u>around a "mean"</u> of natural rate unemployment.

7. Conclusion

A simple aggregated-over-micro-disequilibria model was formulated and tested with results quite favourable and comparable to similar studies. This model predicted that (apart from a correction term and a random error) aggregate employment in the market in question would be given by a convex combination of demand and supply, the weighting factor being the proportion of sectors that are in excess supply. The estimates were close to ones we obtained from the standard disequilibrium switching model, thus justifying the close relation between the two models that we stress throughout the paper. Important caveats involve the very strong positive serial correlation followed by the main equation error, and the questionable predeterminedness of the real (but not nominal) wage.

The signs of estimated coefficients are in general as predicted by theory -- significantly negative effects of wage on demand and of a linear trend (technology, etc.) on supply, significantly positive elasticities of demand with respect to real output and of supply with respect to potential labour force and the real interest rate. A measure of expected change in wage was found to introduce strong collinearity, thus preventing reliable tests of standard predictions of the intertemporal labour substitution hypothesis. In agreement with similar studies, we find a statistically insignificant and close to zero elasticity of supply with respect to the (after tax) wage. The "fiscal illusion" hypothesis on the part of workers was tested and rejected. In general, we found the supply side more satisfactorily determined than in past disequilibrium studies. Our simple model was able to predict quite closely actual employment (within sample) and to give estimates of the extent of disequilibrium in the labour market that correlated closely with intuitive measures of the degree of "slackness" in the market.

For safer inferences on the equilibrium versus disequilibrium issue, however, evidence of a much more disaggregated nature is needed, as non-market clearing makes more sense as a micro phenomenon (see Bouissou *et al.* (1986)). Studies that use very disaggregated data and try to address the prices and quantities flexibility issue include Kawasaki, McMillan and Zimmerman (1982, 1983), Nerlove (1983) and Koenig and Nerlove (1984). The reason that such micro-data based studies are more suited to studying the equilibrium versus disequilibrium question is that once one starts aggregating, the perpetual flux of micro disequilibria that might exist could become extremely hard to detect.

Appendix A

Derivation of Equation (9) From First Principles

Recall the model

(1) $$D_t = D_t^* + \varepsilon_t$$

(2) $$S_t = S_t^* + u_t$$

(3) $$Q_t = \min(D_t, S_t) \ .$$

Excess D ≡ D − S , {excess S} iff $\{D^* - S^* < u - \varepsilon \equiv \eta\}$. Defining $K \equiv D^* - S^*$, $E(\varepsilon | \text{excess S}) = E(\varepsilon | K < \eta)$. Now

(A1) $$P(\varepsilon | K < \eta) \equiv \frac{P(\varepsilon \cap K<\eta)}{P(K<\eta)} = \frac{\int_K^\infty f(\varepsilon,\eta)d\eta}{P(K<\eta)} = \frac{\int_K^\infty f(\varepsilon,\eta)d\eta}{\Phi(-\frac{K}{\sigma_\eta})} \ .$$

Hence

(A2) $$E(\varepsilon | \text{excess S}) = \frac{\int_{-\infty}^\infty \int_K^\infty \varepsilon \ f(\varepsilon,\eta)d\eta d\varepsilon}{\Phi(-\frac{K}{\sigma_\eta})} = \frac{\int_K^\infty [\int_{-\infty}^\infty \varepsilon \ f(\varepsilon|\eta)d\varepsilon]f(\eta)d\eta}{\Phi(-\frac{K}{\sigma_\eta})}$$

Since $\binom{\varepsilon}{\eta} \sim N\{\binom{0}{0}, \begin{pmatrix} \sigma_\varepsilon^2 & \sigma_{\varepsilon\eta} \\ \sigma_{\varepsilon\eta} & \sigma_\eta^2 \end{pmatrix}\}$, we know that

(A3) $$E(\varepsilon | \eta) = 0 + \frac{\sigma_{\varepsilon\eta}\eta}{\sigma_\eta^2} \ .$$

Using A3 in A2 yields:

(A4) $$E(\varepsilon | \text{excess S}) = \frac{\sigma_{\varepsilon\eta}}{\sigma_\eta^2} \frac{1}{\Phi(\frac{-K}{\sigma_\eta})} \int_K^\infty \eta \ f(\eta)d\eta = \frac{\sigma_{\varepsilon\eta}}{\sigma_\eta} \frac{\phi(\frac{K}{\sigma_\eta})}{1-\Phi(\frac{K}{\sigma_\eta})}$$

Result A4 is obtained by noting that

$$\int_K^\infty \frac{\eta \ f(\eta)d\eta}{\sigma_\eta} = \left[-\frac{1}{\sqrt{2\pi}\,\sigma_\eta} \exp\left(-\frac{1}{2}\frac{\eta^2}{\sigma_\eta^2}\right)\right]_K^\infty = \frac{1}{\sqrt{2\pi}\,\sigma_\eta} \exp\left(-\frac{1}{2}\frac{\eta^2}{\sigma_\eta^2}\right) \equiv \phi\left(\frac{K}{\sigma_\eta}\right) \ .$$

Reversing the argument we obtain

$$\text{(A5)} \qquad E(u|\text{excess } D) = \frac{-\sigma_{u\eta}}{\sigma_\eta} \frac{\phi(\frac{K}{\sigma_\eta})}{\Phi(\frac{K}{\sigma_\eta})}$$

Finally we substitute A4 and A5 into equation (5) to get equation (9).

Appendix B

Obtaining Heteroskedasticity Consistent Standard Errors in Eq(9)

The White (1980) - type consistent covariance correction proceeds as follows:

Write EQ = $f(X;\beta)$. Form the vector of derivatives $\frac{\partial f(x;\beta)}{\partial \beta}$. Let s be the estimated standard error of the regression printed by the NLLS program and b be the consistent NLLS estimates. Let V_t be the fitted residuals and create the matrix

$$W = \sum_t V_t^2 \frac{\partial f(x;b)}{\partial \beta} \frac{\partial f(x;b)'}{\partial \beta} .$$

Then a consistent estimate of the covariance matrix for NLLS is $V = s^2PWP$, where P is the (inconsistent) covariance matrix printed by the NLLS program which neglects heteroskedasticity. (See McFadden and Newey (1982) for a proof).

Appendix C
Data Sources and Description of Data Series

Sources: Citibank Economic Data Base, Council of Economic Advisors (CEA), U.S. Board of Governors of the Federal Reserve System

A. Data Series* — Description of Data**

Data Series*	Description of Data**
PRODWORM =	Production workers on non-agricultural payrolls: manufacturing (thousands) (LPWM6)
AWKHRPWM =	Average weekly hours of production workers: manufacturing (LPHRM6)
AHEPWM =	Average hourly earnings of production workers: manufacturing ($) (LE6HM)
PPIM =	Producer price index: manufactured goods (1967 = 100) (PWM)
CPIU =	PI-U: all items (1967 = 100) (PZU)
WBAR =	Compensation per hour employees: non-farm business sector (LBPUR) seasonally adjusted
IP =	Quarterly series on industrial production, U.S. Board of Governors of the Federal Reserve System
LFCIV =	Civilian labour force: total (thousands) (LHC6)
UNEM =	Percent unemployed of civilian labour force, total, 16 years and over (LHUR) seasonally adjusted
YBAR =	Potential GNP in real 1972 dollars (from the Council of Economic Advisors)
TB3MONTH =	Interest rate: U.S. Treasury Bills, auction average, 3-month (% per annum) (FYGN3)
GPY =	Personal income, current dollars
GYD =	Disposable personal income, current dollars

B.

Constructed Series	Definition
LE =	log (PRODWORM*AWKHRPWM*52.)
W =	log (AHEPWM/PPIM)
TH =	1. - GYD/GPY
X =	log (AHEPWM*(1.-TH)/CPIU)
POTL =	log (LFCIV)
RRCHAT =	(1.0 + TB3MONTH)/(1.0 + CINFHAT)
RELW =	(AHEPWM/WBAR) , normalized to 1. in 1977.
Y =	log (IP/PPIM)

Notes:

* All series are seasonally unadjusted, unless otherwise stated.

** Original CITIBASE names of variables are given in parentheses.

Footnotes

[1]This independence of shocks across j might appear as restrictive, given, for example the presence of macroeconomic shocks. Recall, however, that we are for the moment conditioning on such economy-wide shocks. Some of these shocks can be captured by the inclusion of economy-wide variables in D_t^* and S_t^*. Below, we will introduce explicitly temporal macroeconomic randomness of a specific nature. In future work, we plan to experiment with more flexible ways of modelling macro shocks.

[2]It is interesting to note the analogy between equation (4') and a model presented in Spencer (1975), also aimed at meeting the criticism of the implausibility of discrete switching between regimes. Spencer proposed, as Quandt (1982) notes "without any obvious economic justification":

$$Q_t = (1-\pi_t)\, D_t + \pi_t\, S_t + \eta_t \tag{F1}$$

where π_t is the cumulative logistic

$$\pi_t = \exp((D_t-S_t)/\theta) \,/\, (1-\exp((D_t-S_t)/\theta)).$$

As shown in Appendix 2 of Hajivassiliou (1983), this follows exactly from equation (4), if we make the assumption that ε and η follow identical and independent <u>extreme value</u> distributions. The problem with (F1) is that if indeed Spencer had this kind of rationalization in mind in justifying why Q should be a convex combination of D and S, then η should <u>not</u> be a Gauss-Markov error as implied by ε^*, u^*.

[3]This requirement is admittedly strong, as the unobservable macroeconomic shocks are assumed not to affect differentially the two sides.

[4]Heteroskedasticity arises because v is a mixture of errors that are conditioned on whether event A or B occurs.

[5]The main idea of smoothing-by-aggregation in disequilibrium models was also pursued empirically in a recent paper by Andrews and Nickell (1986).

[6]For an equivalent final equation in the case of the errors being i.i.d. <u>extreme-value</u>, see Appendix 2 of Hajivassiliou (1983). An interesting generalization of the model is to recognize that θ and η are not perfectly correlated. In such a case, eq(9) reads

$$Q = D^*\cdot(1 - \Phi(z)) + S^*\cdot\Phi(z) - \sigma\, \phi(z) + \Phi(z)\cdot\eta + (1 - \Phi(z))\cdot\theta \;, \tag{F2}$$

where $z \equiv (D^* + \theta - S^* - \eta)/\sigma$, and (η, θ) is jointly normally distributed. This yields an implicit non-linear model with non-additive errors, which may be able to explain at least part of the serial correlation we find below, because of the presence of the last two terms. This approach will be pursued in future work.

[7]Quandt (1986) presents Monte-Carlo evidence that this correction for heteroskedasticity is well warranted. One consequence of the heteroskedasticity is that the standard errors printed by a normal NLLS package will be inconsistent, thus leading to wrong inferences. The solution to this problem is described in Appendix B. A second property of NLLS applied on eq. (5') is that the resulting estimates will be less efficient that maximum likelihood estimates of the switching model because the error v is both heteroskedastic and a mixture of truncated errors of ε and u. Applying the appropriate heteroskedasticity weighting through weighted NLLS would only raise enormously the computational burden involved, without eliminating the second cause of inefficiency in our case. (For a model in a similar vein where explicit treatment of heteroskedasticity is worked out see Dubin and McFadden (1984)).

[8]Some attempts at testing the assumed predeterminedness of the wage using Hausman (1978) tests in the context of equilibrium models, are reported in Hajivassiliou (1983) with mixed results: (W = real wage facing employers, X = after tax real wage facing workers, Y = real industrial production)

a. Test exogeneity of X and W, assuming Y exogenous:

   LR = 10.8, $\chi^2(2)$ 1% = 9.21, $\chi^2(2)$ 5% = 5.99

b. Test exogeneity of W and Y

   LR = 13.95, $\chi^2(2)$ 1% = 9.21, $\chi^2(2)$ 5% = 5.99

c. Test exogeneity of X, assuming Y exogenous:

   LR = 8.95, $\chi^2(1)$ 1% = 6.64, $\chi^2(1)$ 5% = 3.84

These rejections are not very strong at the 1% level of significance, so the evidence against adopting the assumption of predetermined wages is weak.

[9]Recall that expectations applied on (3) in order to get the final equation (9) were taken conditional on X_t^d, X_t^s as predetermined variables. If such exogeneity assumptions are violated, the particular functional form of eq (9) changes. Hence NLLS estimation of eq (9) as derived in the text would yield inconsistent estimates. For this reason instrumental variables techniques like two-stage NLLS would also not be able to solve the problem. The solution is to specify a price adjustment rule and work through the expected quantity equation using arguments precisely analogous to the ones used in Appendix A, thus taking full account of the endogeneity of the wage (through the jointly normal distribution of ε, u and ξ). The exact form of equation (9) has been derived by the author in the explicit case of adjustment rule (10). The equation is again of the general form

$$(9'') \qquad Q = (1-P) \cdot D^* + P \cdot S^* - \text{correction term}$$

and is not reproduced here. The exact expression for (9") depends fundamentally on (10); however, specification of (10) is highly likely to be precarious given that, for example, unionization rates have been very steadily falling in the U.S. economy. Fairly strong evidence on the problematic nature of (10) is contained in Quandt (1978) and Romer (1981) where slight perturbations of the wage adjustment specification gave quite erratic behavior of the estimates.

[10]Validity of inferences would be affected though, as the standard errors would be inconsistent. No correction for such possibility was attempted. On this issue, see Newey and West (1987).

[11]In a previous version of this paper that worked with seasonally adjusted data, we found evidence of very strong (positive) serial correlation that suggested the use of first differences of (log) variables as the most adequate empirical approximation. In view of the noted non-linearities of the model, however, the operation of differencing is not innocuous. The easiest way to see this is through eq (3): because the min $(\cdot,\cdot)$ operation is nonlinear and also not one-to-one,

$$\Delta \min(D,S) \neq \min(\Delta D, \Delta S).$$

Moreover, the economic interpretation of the model is now complicated since it is not at all clear why the change in quantity should be given by the minimum of the change in demand and the change in supply. This is exactly true only if the previous period was one of equilibrium (zero excess demand). There are, however, circumstances under which the differencing procedure is appropriate. Suppose that agents recognize the possibility of being rationed in the coming period. Suppose further that, in forming their notional demand for labour, the employers look at the level of actual employment last period in relation to their notional demand of last period. This would provide a measure of how constrained they were at t-1 and through extrapolative expectations, a notion of how constrained they expect to be this period. Anticipating a positive probability of rationing would therefore be expected to depress the notional magnitudes. (For a model where the probability of rationing affects the notional supply of labor, see Eaton and Quandt (1983)). Depressing effects of a positive probability of rationing on the supply side are analogous to the so-called "discouraged worker" phenomenon. Hence we obtain

$$(1'') \qquad L_t^d = \gamma(L_{t-1} - X_{t-1}^d\beta^d) + X_t^d\beta^d + \epsilon_t - \gamma\epsilon_{t-1}.$$

and for exactly analogous reasons

$$(2'') \qquad L_t^s = \delta(L_{t-1} - X_{t-1}^s\beta^s) + X_t^s\beta^s + u_t - \delta u_{t-1}.$$

Working with first differences then amounts to assuming that (a) $\gamma=\delta=1$ and (b) ϵ and u are random walks, so that $\epsilon-\gamma\epsilon_{-1}$ and $u-\delta u_{-1}$ are Gauss-Markov errors. In that case,

$$(3'') \qquad \Delta L_t = \min(\Delta(X_t^d\beta^d + \epsilon_t), \Delta(X_t^s\beta^s + u_t)),$$

which is exactly what the differencing formulation corresponds to. In our empirical results we found that the $\gamma=\delta=1$ hypothesis was not rejected by the data and that after differencing no evidence of serial correlation remained. The test of the $\gamma = \delta = 1$ hypothesis is a straightforward Lagrange multiplier test against the alternative of γ, δ being freely estimated. Even the equivalent (locally) Wald Statistic, which usually rejects the most frequently of the "trinity" of tests, failed to reject with a χ^2 value of only 0.01. Use of the Lee (1984a) DW-like procedure of testing for AR(1) in disequilibrium models did not point out to any remaining (AR(1)) serial dependence, once first differenced deseasonalized data were used.

[12]Bodkin (1969) and Geary and Kennan (1982) find significant differences in results depending on whether the CPI or WPI is used. To pursue this question, we also estimate (but do not present here) versions that use the CPI to deflate both the product wage and the worker wage, as most studies do (see e.g., Neftci (1978)). Theory appears to be confirmed as the fit of equation (9) worsens uniformly, once this theoretically inappropriate deflating is used. Interestingly enough, the estimated parameter values do not move greatly, despite the noted deterioration in the fit: unlike the Geary and Kennan (1982) study, we do not find any spectacular differences depending on the method of deflating the product wage.

[13]A more appropriate approach would be to model explicitly spillovers between labour and goods markets.

[14]Though the standard choice-at-the-goods- leisure-margin model predicts a role for unearned income (for fixing the intercept of the budget constraint), no such term was included. The point first made by Romer (1981) and accepted by Quandt (1981) is that the nonsensical results obtained by Rosen and Quandt (1978) (for example, all the years of the Great Depression were predicted to be in excess demand by the model!) were mainly caused by the behavior of the unearned income term, in that through the basic life cycle model unearned income is endogenous and positively associated with labour supply. Instead of attempting the intractable task of modelling the lag structure from the life-cycle model that relates unearned income and hours of work, dropping unearned income from the model altogether seems to be a reasonable way to proceed.

[15]"LR" is not exactly the usual likelihood ratio test, given that estimation of eq (9) is by NLLS. It is a valid measure of "distance", however, and also distributed asymptotically as χ^2 (d.f.) Moreover, there is evidence that it possesses high power, particularly when used as a specification test.

[16]Unfortunately this series is seasonally adjusted.

References

Altonji, J.G. (1982): "The Intertemporal Substitution Model of Labour Market Fluctuations: An Empirical Analysis," Review of Economic Studies, XLIX.

Altonji, J.G. and O. Ashenfelter (1980): "Wage Movements and the Labour Market Equilibrium Hypothesis," Economica, (August).

Andrews, M. and S. Nickell (1986): "A Disaggregated Disequilibrium Model of the Labour Market," Oxford Economic Papers, v.38, 3 (November).

Barro, R.J. (1977): "Unanticipated Money Growth and Unemployment in the U.S.," American Economic Review, 67 (March).

Barro, R.J. (1978): "Unanticipated Money, Output and the Price Level in the U.S.," Journal of Political Economy, 86 (August).

Barro, R.J. and H.I. Grossman (1971): "A General Disequilibrium Model of Income and Employment," American Economic Review, 61 (March).

Bodkin, R.J. (1969): "Real Wages and Cyclical Variations in Employment: A Reexamination of the Evidence," Canadian Journal of Economics, 2 (August).

Bouissou, M.B., J.J. Laffont and Q.H. Vuong (1986): "Disequilibrium Econometrics on Micro Data," Review of Economic Studies, LIII.

Bowden, R.J. (1978): "Specification, Estimation and Inference for Models of Markets in Disequilibrium," International Economic Review, 19 (October).

Box, G.E.P. and G.M. Jenkins (1970): Time Series Analysis, Forecasting and Control. San Francisco, Holden Day.

CITIBASE: Citibank Economic Database, New York, Citibank. 1985 Update.

Dubin, J.A. and D. McFadden (1984): "An Econometric Analysis of Residential Electric Appliance Holdings and Consumption," Econometrica, 52 (March).

Eaton, J. and R.E. Quandt (1983): "A Quasi-Walrasian Model of Rationing and Labor Supply," Economica (August).

Engle, R. Hendry, D.F. and J.F. Richard (1983): "Exogeneity," Econometrica, January.

Fair, R.C. and D.M. Jaffee (1972): "Methods of Estimation for Markets in Disequilibrium," Econometrica, 40 (May).

Geary, P.T. and J. Kennan (1982): "The Employment-Real Wage Relationship: An International Study," Journal of Political Economy, 90 (August).

Godley, W.A.H. and W.D. Nordhaus (1972): "Pricing in the Trade Cycle," Economic Journal, 82 (September).

Goldfeld, S.M. and R.D. Quandt (1975): "Estimation of a Disequilibrium Model and the Value of Information," Journal of Econometrics, 3.

Goldfeld, S.M. and R.E. Quandt (1978): "Some Properties of the Simple Disequilibrium Model with Covariance," Economics Letters, 1,4.

Goldfeld, S.M. and R.E. Quandt (1979): "Estimation in Multi-market Disequilibrium Models," Economics Letters, 4,4.

Granger, C.W.J. and P. Newbold (1977): Forecasting Economic Time Series, New York, Academic Press.

Hajivassiliou, V.A. (1983): "Estimating Models of the U.S. Aggregate Labour Market and Testing for Disequilibrium," MIT mimeo.

Hajivassiliou, V.A. (1985): "The Estimation of Certain Non-Linear Models with Autocorrelated Errors by Conditional Quasi Maximum Likelihood Methods," MIT mimeo.

Hajivassiliou, V.A. (1986a): "Temporal Dependence in Limited Dependent Variables Models: Theoretical and Monte Carlo Results," Cowles Foundation Discussion Paper #803.

Hajivassiliou, V.A. (1986b): "Two Misspecification Tests for the Simple Switching Regressions Disequilibrium Model," Economics Letters, 22.

Hansen, L.P. and Singleton (1982): "Generalized Instrumental Variables Estimation of Nonlinear Rational Expectations Models," Econometrica, 50 (September).

Harvey, A.C. (1981): Time Series Models. John Wiley and Sons, New York.

Hausman, J.A. (1978): "Specification Tests in Econometrics," Econometrica, 46 (November).

Hausman, J. A. (1983): "Specification and Estimation in Simultaneous Equations Models," in Handbook of Econometrics, Vol.1, Griliches,Z. and M. Intrilligator (eds.). North-Holland: Amsterdam.

Heckman, J.J. (1979): "Sample Selection Bias as a Specification Error," Econometrica, 47 (January).

Johnson, N. and S. Kotz (1972): Distributions in Statistics: Continuous Multivariate Distributions. New York: John Wiley and Sons.

Kawasaki, S. McMillan, J., and K.F. Zimmerman (1982): "Disequilibrium Dynamics: An Empirical Study," American Economic Review, 72.

Kawasaki, S., McMillan, J. and K.F. Zimmerman (1983): "Inventories and Price Inflexibility," Econometrica, 51 (May).

Koenig, H. and M. Nerlove (1984): "Response of Prices and Production to Unanticipated Demand Shocks: Some Microeconomic Evidence," mimeo.

Kooiman, P. (1984): "Smoothing the Aggregate Fix-price Model and the Use of Business Survey Data," The Economic Journal, December.

Laffont, J.J. and A. Monfort (1979): "Disequilibrium Econometrics in Dynamic Models," Journal of Econometrics, 11.

Lambert, J.P. (1984): "Models Macro-Economique de Rationnement et Enquettes de Conjoctures," Les Recherches Economiques de Louvain: Louvan-La-Neuve, Belgium.

Lee, L.-F. (1984a): "The Likelihood Function and a Test for Serial Correlation in a Disequilibrium Market Model," Economics Letters, V. 14, Nos. 2-3.

Lee, L.-F. (1984b): "Regime Classifications in the Disequilibrium Market Models," Economics Letters, V. 14, Nos. 2-3.

Maddala, G. and F. Nelson (1974): "Maximum Likelihood Methods for Markets in Disequilibrium," Econometrica, 42.

Malinvaud, E. (1980): "Macroeconomic Rationing of Employment," in E. Malinvaud and J.P. Fitoussi, Unemployment in Western Countries, London, MacMillan.

Mankiw, N.G., J. Rotemberg and L. Summers (1985): "The Intertemporal Hypothesis in Macroeconomics," Quarterly Journal of Economics (February).

McFadden, D. and W. Newey (1982): "Asymptotic Properties of Non-Linear Estimators," MIT mimeo.

Monfort, A. (1982): Discussion of Quandt's (1982) "Survey of Disequilibrium Models," Econometric Reviews, 1.

Muellbauer, J. (1978): "Macro Theory vs. Macro Econometrics: The Treatment of Disequilibrium in Macro Models," Birkbeck College Discussion Paper, 79 (April).

Neftci, S. (1978): "A Time Seris Analysis of the Real Wages - Employment Relationship," Journal of Political Economy (April).

Nerlove, M. (1983): "Expectations, Plans and Realizations in Theory and Practice," Econometrica, 51 (September).

Newey, W. and K. West (1987): "A Simple, Positive Definite, Heteroskedasticity and Autocorrelation Consistent Covariance Matrix," forthcoming in Econometrica.

Nordhaus, W. (1976): "Recent Developments in Price Dynamics," in O. Eckstein (ed.), The Econometrics of Price Determination (S.S.R.C./Fed).

Quandt, R.E. (1981): "Autocorrelated Errors in Simple Disequilibrium Models," Economics Letters, 7.

Quandt, R.E. (1982): "Econometric Disequilibrium Models," Econometric Reviews, 1.

Quandt, R.E. (1986): "Estimation in Disequilibrium Models with Aggregation," Financial Research Center Memorandum No. 68, Princeton University.

Robinson, P. (1982): "On the Asymptotic Properties of Models Containing Limited Dependent Variables," Econometrica (January).

Romer, D. (1981): "Rosen and Quandt's Disequilibrium Model of the Labor Market: A Revision," Review of Economics and Statistics.

Rosen, H.S. and R.E. Quandt (1978): "Estimation of a Disequilibrium Aggregate Labor Market," Review of Economics and Statistics, 60.

Sargent, T.J. (1978): "Estimation of Dynamic Labour Demand Schedules Under Rational Expectations," Journal of Political Economy (December).

Spencer, P.D. (1975): "A Disequilibrium Model of Personal Sector Bank Advances 1964-1974: Some Preliminary Results," Working Paper, H.M. Treasury.

Sutton, J. (1981): "A Formal Model of the Long-Run Phillips Curve," Economica (November).

Tobin, J. (1972): "Inflation and Unemployment," American Economic Review (March).

White, H. (1980): "A Heteroskedasticity-Consistent Covariance Matrix and a Direct Test for Heteroskedasticity," Econometrica, 48.

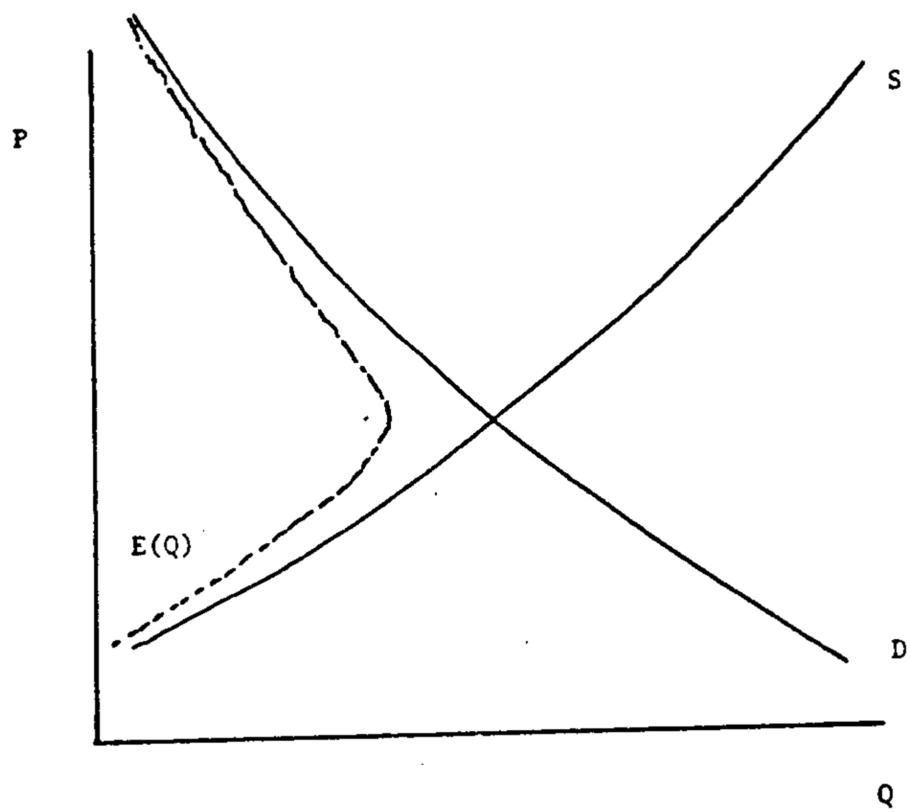
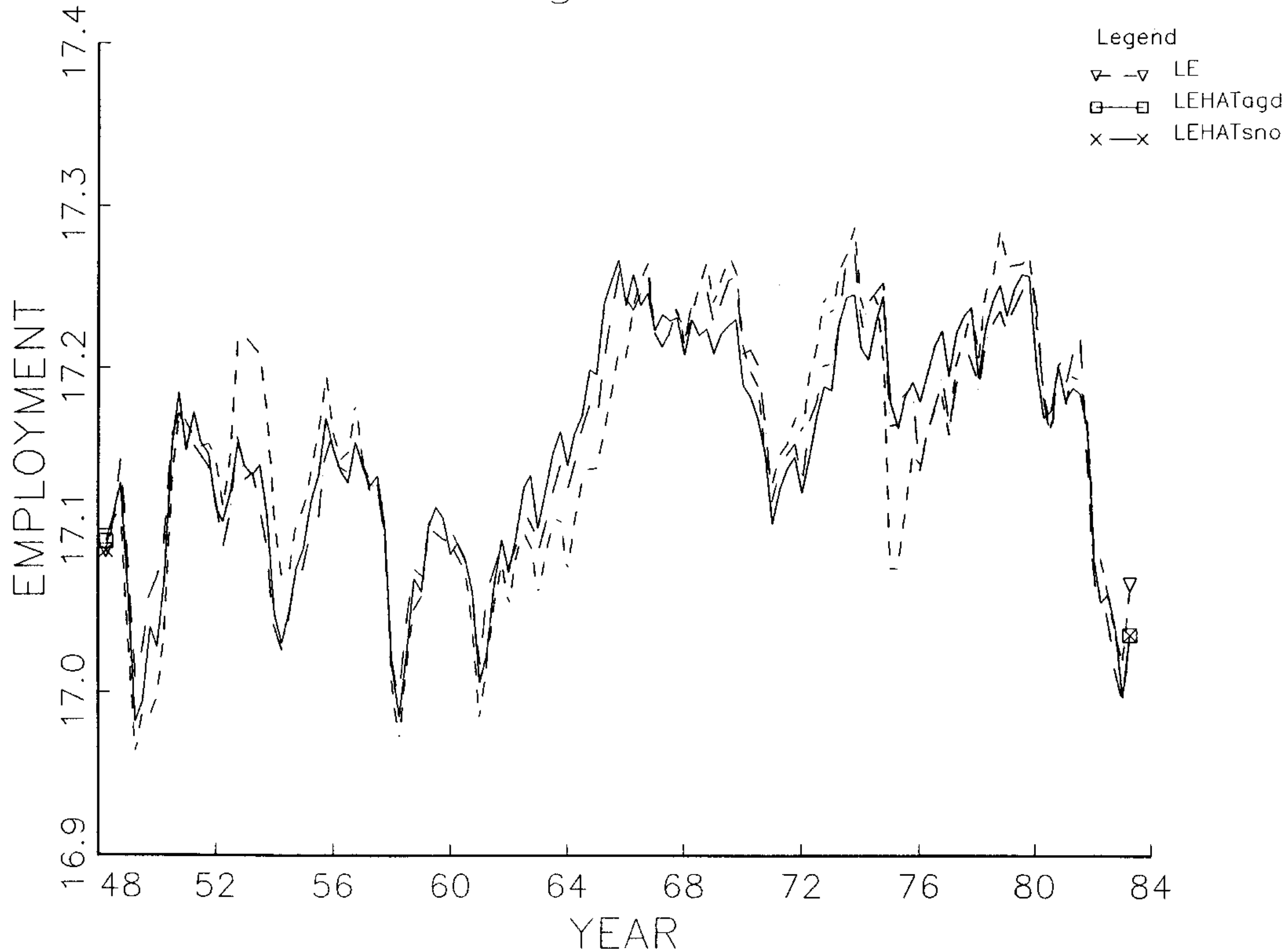


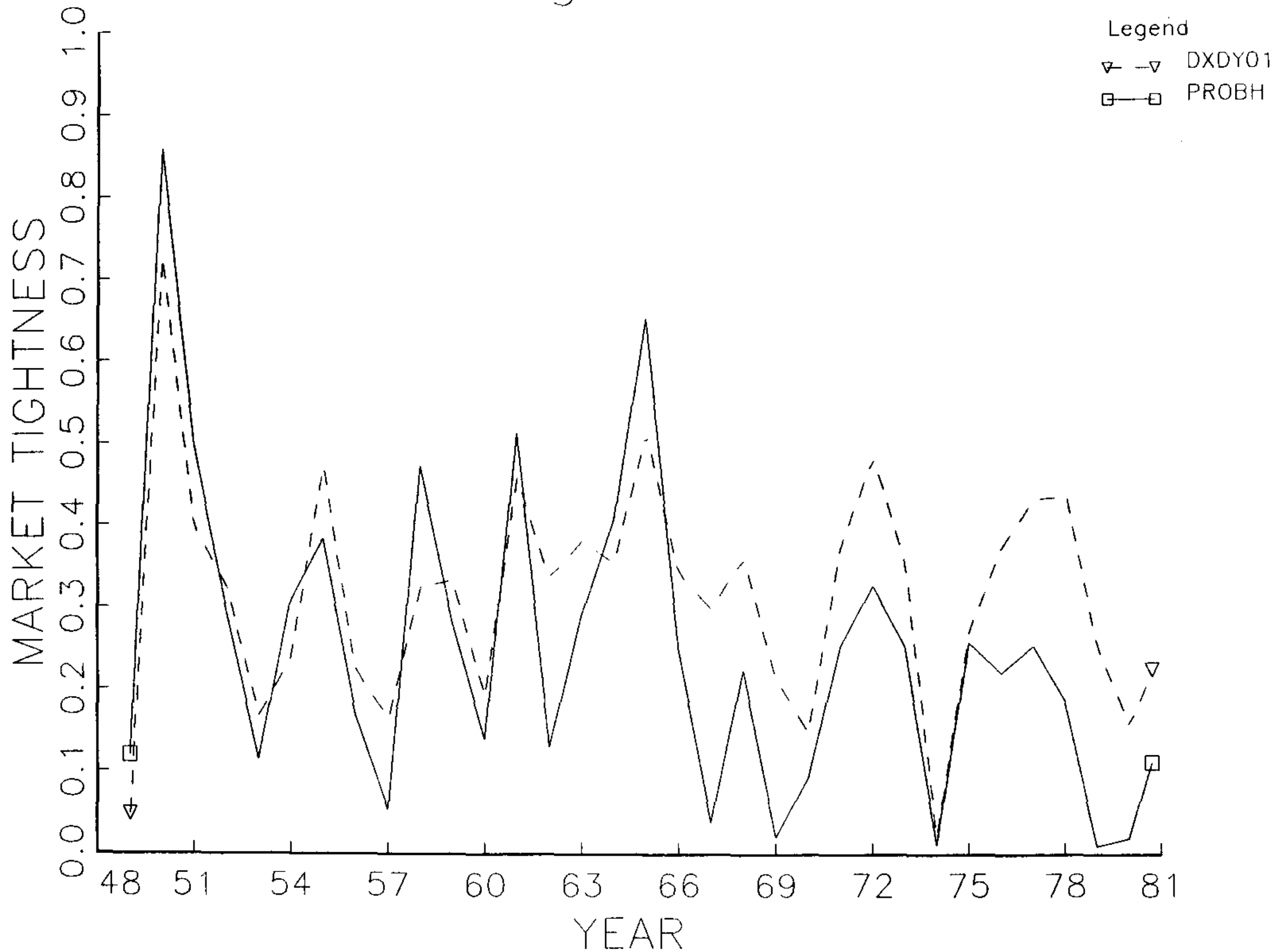
Figure 1

Figure 2



Legend
LE
LEHATagd
LEHATsno
EMPLOYMENT
17.4
17.3
17.2
17.1
17.0
16.9
48
52
56
60
64
68
72
76
80
84
YEAR

Figure 3



Legend
DXDY01
PROBH
MARKET TIGHTNESS
0.0
0.1
0.2
0.3
0.4
0.5
0.6
0.7
0.8
0.9
1.0
48
51
54
57
60
63
66
69
72
75
78
81
YEAR

TABLE 1

Forecasting equations for the inflation rate (CINF), the real product wage (W) and the post-tax real wage (X) (t statistics in parentheses)

A. The preferred equations used in the estimations.

(1) $$CINF = \underset{(-2.529)}{-0.354E{-}02} + \{seasonals\} + \underset{(3.342)}{0.522E{-}04{*}T} + \underset{(10.519)}{0.667{*}CINF(-1)}$$

$LF = 523.41$ $SSR = 0.4638E{-}02$ $DW = 2.016$ $R^2 = 0.680$ $T = 141$

(2) $$W = \underset{(-1.227)}{-0.524E{-}01} + \{seasonals\} + \underset{(0.562)}{0.272E{-}04{*}T} + \underset{(19.212)}{1.463{*}W(-1)} - \underset{(-6.304)}{0.479{*}W(-2)}$$

$LF = 448.51$ $SSR = 0.1351E{-}01$ $DW = 2.049$ $R^2 = 0.997$ $T = 141$

(3) $$X = \underset{(-0.786)}{-0.579E{-}01} + \{seasonals\} + \underset{(0.631E-01)}{0.443E{-}05{*}T} + \underset{(55.705)}{0.982{*}X(-1)}$$

$LF = 448.75$ $SSR = 0.1353E{-}01$ $DW = 1.993$ $R^2 = 0.996$ $T = 141$

B. Some alternative specifications showing why (1),(2),(3) preferred.

(4) $$CINF = \underset{(-2.509)}{-0.353E{-}02} + \{seasonals\} + \underset{(3.151)}{0.513E{-}04{*}T} + \underset{(7.552)}{0.655{*}CINF(-1)} + \underset{(0.206)}{0.018{*}CINF(-2)}$$

$LF = 523.43$ $SSR = 0.464E{-}02$ $DW = 1.999$ $R^2 = 0.674$ $T = 141$

(5) $$W = \underset{(-0.689)}{-0.333E{-}01} + \{seasonals\} - \underset{(0.494)}{0.266E{-}04{*}T} + \underset{(81.840)}{0.988{*}W(-1)}$$

$LF = 430.20$ $SSR = 0.176E{-}01$ $DW = 1.047$ $R^2 = 0.996$ $T = 141$

(6) $$X = \underset{(-0.763)}{-0.568E{-}01} + \{seasonals\} + \underset{(0.407E-01)}{0.291E{-}05{*}T} + \underset{(11.182)}{0.972{*}X(-1)} + \underset{(0.121)}{0.106E{-}01{*}X(-2)}$$

$LF = 448.76$ $SSR = 0.135E{-}01$ $DW = 1.976$ $R^2 = 0.996$ $T = 141$

TABLE 2

Equation (9), T = 141, Dependent Variable = LE

(Asymptotic t-ratios in parentheses, heteroskedasticity-consistent for the AGD models)

Model	AGD1	AGD2	SNOR1	SNOR2
LF	285.273	275.686	316.54	287.915
SSR	0.1392	0.1597	---	---
CORR(LE,LEFIT)	0.951	0.932	0.916	0.908
<u>Demand</u>				
Constant	5.451	6.253	9.298	10.333
	(4.979)	(5.139)	(11.510)	(15.119)
Q1	1.197	1.110	-0.018	-0.025
	(10.851)	(9.070)	(-1.655)	(-2.196)
Q2	-0.026	-0.026	-0.019	-0.031
	(-1.947)	(-1.973)	(-1.662)	(-2.783)
Q3	-0.031	-0.032	-0.008	-0.015
	(-2.501)	(-2.457)	(-0.067)	(-1.249)
Trend	0.007	0.109	0.008	0.007
	(5.381)	(7.812)	(10.531)	(11.03)
W	-2.133	-2.261	-1.435	-1.907
	(-3.127)	(-3.279)	(-3.012)	(-4.176)
WLAG	-0.915	-0.581	-0.612	0.128
	(-1.193)	(-0.755)	(-1.211)	(0.281)
Y	0.861	0.851	0.735	0.644
	(5.973)	(5.284)	(6.057)	(5.679)
YLAG	0.344	0.291	0.131	0.123
	(2.546)	(1.920)	(1.279)	(1.234)
SIGD	---	---	0.041	0.042
			(11.893)	(12.383)
<u>Supply</u>				
Constant	20.972	18.699	19.983	18.242
	(17.680)	(13.341)	(26.845)	(21.265)
Q1	-0.563	-0.088	-0.038	-0.034
	(-2.070)	(-0.281)	(-1.204)	(-4.562)
Q2	-0.019	-0.031	0.053	-0.059
	(-1.338)	(-1.449)	(0.611)	(-2.209)
Q3	-0.011	-0.044	0.011	-0.043
	(-0.336)	(-1.062)	(1.319)	(-1.764)

(continued)

T	-0.018	-0.038	-0.031	0.012
	(-0.584)	(-0.965)	(-0.474)	(1.182)
X	1.293	1.137	0.249	-0.621
	(1.930)	(1.357)	(1.180)	(-1.285)
XLAG	0.342	0.190	1.193	1.351
	(0.612)	(0.262)	(6.425)	(2.369)
RELW	1.689	1.108	1.663	1.653
	(1.051)	(0.616)	(4.811)	(2.053)
RELWLAG	0.960	2.347	0.994	0.050
	(0.639)	(1.302)	(2.828)	(-0.064)
POTL	1.087	1.558	0.757	1.385
	(0.999)	(1.123)	(2.765)	(1.356)
POTLLAG	-0.582	-1.292	-0.880	-1.643
	(-0.530)	(-0.930)	(3.162)	(-1.561)
RRCHAT	0.068	---	0.082	---
	(6.781)		(15.409)	
SIGS	---	---	0.004	0.005
			(5.030)	(3.016)
SIGH	0.101	0.133	---	---
	(3.656)	(3.534)		
RHO(D,S)	---	---	0.093	0.112
			(0.473)	(0.568)

TABLE 3

Equation (9), T = 141, Dependent Variable = LE

Serial Correlation and Wage-Exogeneity Tests

(Asymptotic t-ratios in parentheses, heteroskedasticity-consistent for AGD1)

Model	AGD1	AGD1SC	SNOR1	SNOR1RW	SNOR1NW
LF	285.273	356.71	316.54	323.94	316.63
SSR	0.1392	0.0262	---	---	---
CORR(LE,LEFIT)	0.951	0.983	0.916	---	---
Demand					
Constant	5.451	6.407	9.298	9.714	10.263
	(4.979)	(2.071)	(11.510)	(14.041)	(13.346)
Q1	1.197	1.113	-0.018	-0.027	-0.025
	(10.851)	(-0.813)	(-1.655)	(-2.610)	(-2.160)
Q2	-0.026	-0.016	-0.019	-0.039	-0.031
	(-1.947)	(-1.325)	(-1.662)	(-3.236)	(-2.834)
Q3	-0.031	-0.024	-0.008	-0.019	-0.015
	(-2.501)	(-1.123)	(-0.067)	(-1.677)	(-1.248)
Trend	0.007	0.011	0.008	0.007	0.007
	(5.381)	(3.646)	(10.531)	(11.566)	(9.828)
W	-2.133	-2.121	-1.435	-3.758	-1.931
	(-3.127)	(-2.161)	(-3.012)	(-4.313)	(-4.134)
WLAG	-0.915	-0.674	-0.612	-1.815	-0.134
	(-1.193)	(-0.678)	(-1.211)	(-2.094)	(-0.274)
Y	0.861	1.125	0.735	1.071	0.652
	(5.973)	(4.312)	(6.057)	(7.328)	(5.629)
YLAG	0.344	0.009	0.131	0.259	0.121
	(2.546)	(1.374)	(1.279)	(1.904)	(1.177)
WRES	---	---	---	3.993	---
				(3.054)	
XRES	---	---	---	2.036	---
				(3.036)	
WNOMRES	---	---	---	---	-0.410
					(-0.384)
SIGD	---	---	0.041	0.038	0.042
			(11.893)	(12.423)	(12.329)

(continued)

Supply

Constant	20.972	18.505	19.983	19.724	18.152
	(17.680)	(7.602)	(26.845)	(24.470)	(37.679)
Q1	-0.563	-0.039	-0.038	-0.040	-0.034
	(-2.070)	(-1.271)	(-1.204)	(-7.900)	(-4.693)
Q2	-0.019	-0.053	0.053	-0.004	-0.059
	(-1.338)	(-0.965)	(0.611)	(-0.315)	(-2.756)
Q3	-0.011	-0.041	0.011	0.006	-0.043
	(-0.336)	(-0.712)	(1.319)	(0.528)	(-2.132)
T	-0.018	-0.032	-0.031	-0.027	-0.014
	(-0.584)	(-0.362)	(-0.474)	(-0.351)	(-1.227)
X	1.293	0.987	0.249	-0.561	-0.628
	(1.930)	(1.606)	(1.180)	(-0.705)	(-1.417)
XLAG	0.342	0.352	1.193	1.989	1.360
	(0.612)	(0.581)	(6.425)	(2.679)	(2.771)
RELW	1.689	1.415	1.663	1.779	1.719
	(1.051)	(0.970)	(4.811)	(4.569)	(2.315)
RELWLAG	0.960	2.252	0.994	1.039	-0.042
	(0.639)	(0.873)	(2.828)	(2.728)	(-0.054)
POTL	1.087	1.281	0.757	0.755	1.391
	(0.999)	(0.691)	(2.765)	(2.658)	(1.732)
POTLLAG	-0.582	-1.083	-0.880	-0.837	-1.678
	(-0.530)	(-0.567)	-(3.162)	(-2.894)	(-1.983)
RRCHAT	0.068	0.079	0.082	0.084	0.087
	(6.781)	(7.632)	(15.409)	(14.957)	(14.307)
WRES	---	---	---	0.553	---
				(1.067)	
XRES	---	---	---	1.103	---
				(1.113)	
WNOMRES	---	---	---	---	0.208
					(0.357)
SIGS	---	---	0.004	0.004	0.005
			(5.030)	(4.006)	(2.682)
SIGH	0.101	0.157	---	---	---
	(3.656)	(4.519)			
RHO(V,VLAG)	---	0.873	---	---	---
		(8.909)			
RHO(D,S)	---	---	0.093	0.082	0.076
			(0.473)	(0.317)	(0.412)

TABLE 4

	PROBH	PROBHC	PHAT	XDY	U
PROBH	1.000				
PROBHC	0.926	1.000			
PHAT	0.922	0.857	1.000		
XDY	0.546	0.480	0.523	1.000	
U	-0.560	-0.510	-0.550	-0.825	1.000