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MACROECONOMIC ANALYSES AND MICROECONOMIC ANALYSES OF LABOR SUPPLY

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Macroeconomic Analyses and Microeconomic Analyses of Labor Supply

ABSTRACT

This paper reports on the current status of the microeconomic research on labor supply behavior. The purpose is to direct attention to microeconomic research that may be helpful in the continuing evaluation of aggregate models designed to explain the dynamic behavior of wages, employment and unemployment. The approach is hopelessly empirical, and the emphasis throughout is on models specified completely enough to allow confrontation with the kind of data actually available.

The first part of the paper is addressed to microeconomists, however. It is a brief attempt to provide a sketch of the stylized facts that aggregate models of the labor market are meant to address. These include (1) the serial 'persistence' in the change in unemployment (or employment), (2) the absence of persistence in the change in the real wage rate, and (3) the continued existence of a negative correlation between nominal price changes and unemployment rates.

The microeconomic (longitudinal) data turn out to be difficult to square up with the simplest life-cycle models of labor supply. Contrary to the predictions of the models, the data indicate that (1) average hours and average real wages move in the same direction only some of the time, and that (2) the within life-cycle, person-specific correlation between hours and wages is negative. The microeconomic (experimental) data indicate other puzzles. More elaborate models incorporating measurement error, non-separable preferences, and unanticipated wage movements may explain these findings, but they are also likely to contain parameters that are not easily identified with the kind of data actually available.

Perhaps an alternative approach may be more fruitful in reconciling the long run determination of hours worked by worker preferences with the short run interaction of observed employment and earnings.

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When it comes to public policy discussions of the labor market, there is no doubt that the big picture of time-series movements in unemployment, employment and wages is the major topic of public concern. Since there is a great deal of active empirical microeconomic research on labor markets, it seems natural to occasionally inquire as to the implications of the findings from this research for the larger issues of macroeconomics. I have often found that this is not always a straightforward undertaking, however, as very little of this microeconomic research is designed with an eye toward its broader implications. This may be all for the best, as it keeps the microeconomists away from the heat of battle and maybe also from the temptations to cook the data. Nevertheless, it is often the case that a line of research will cut across both macroeconoomic and microeconomic issues, and when this does not happen, it may even be a signal that something is amiss in one area or another.

In this paper, I report mainly on my perception of the current status of the microeconomic research on labor supply behavior. The purpose is to direct attention to microeconomic research that may be helpful in the continuing evaluation of aggregate models designed to explain the dynamic behavior of wages, employment, and unemployment. My approach is hopelessly empirical and the emphasis throughout is on models specified completely enough to allow confrontation with the kind of data actually available.

The first part of the paper is addressed to microeconomists, however. It is a brief attempt to provide a sketch of the stylized facts that aggregate models of the labor market are meant to address. The goal here is to summarize for (a perhaps skeptical) reader some simple empirical regularities that seem strong enough to deserve explanation. If they are presented in a convincing enough fashion, perhaps even the microeconomists will find them worthy of attention.

I. Facts and Theories of the Aggregate Labor Market

Although the catalogue could surely be longer, I would like to emphasize three basic characteristics of the aggregate labor market that must surely top the list of important empirical regularities. The first of these is the high serial correlation or "persistence" in measures of the change in unemployment (or employment). The existence of this persistence is relatively easy to document and its existence is rarely questioned. Here I want only to indicate the nature of the serial correlation and indicate its striking similarity in several countries.

The second empirical regularity I should like to emphasize is the absence of persistence in measures of the change in the aggregate real wage rate. Especially for the U.S., the aggregate real wage is very close to a random walk (with drift) and only weakly correlated with nominal variables. The absence of persistence in the change in the real wage is less well known, and perhaps more difficult to document, than is the presence of persistence in the change in the unemployment rate.

The third empirical regularity I should like to emphasize is the consistent existence of a correlation between nominal price changes and

Table 1

Second Order Univariate Autoregressions for

Unemployment, Annual Data; U.S., U.K., Canada*

Country	U.S.	U.K.	Canada	U.S.	U.K.	Canada
Period of Fit	1894 1983	1894 1983	1923 1983	1946 1983	1946 1983	1946 1983
Coefficient of:						
Unemployment Lagged Once (Standard Error)	1.169 (.096)	1.190 (.104)	1.185 (.124)	.589 (.169)	.982 (.165)	.735 (.173)
Unemployment Lagged Twice (Standard Error)	−0.350 (.096)	−0.289 (.107)	−0.347 (.124)	−.164 (.169)	−.521 (.195)	−.071 (.191)

*These regressions also contain quadratic time trends.

Source: (i) Yearbook of Labour Statistics, International Labour Office (1982, and other issues),

(ii) Economic Report of the President 1984,

(iii) Historical Statistics of the United States, from Colonial times to 1970,

(iv) European Historical Statistics 1750–1970.

unemployment rates. The existence of this "Phillips Curve" relationship is sometimes doubted by the skeptics, although the high unemployment/lower inflation experience of the last two years in the U.S., Canada, and Europe must surely have caused even the doubters to think twice. There is nevertheless plenty of room for skepticism about the nature and even the existence of this relationship, so that the question of whether a simple method exists for convincing the doubters may remain open.

A. Persistence in the Change in Unemployment

The first three columns of Table 1 contain the fit of second-order autoregressions to annual data on the unemployment rate for the U.S., the United Kingdom and Canada for various time periods. The results in these tables reveal three relatively straightforward "facts" about time-series movements in aggregate unemployment. First, the unemployment rate in none of these countries can be well represented by a simple first-order autoregressive process. Shocks to unemployment result in the hump-shaped moving average representation reported in columns 1-3 of Table 2 for long time-series of annual data or for quarterly or monthly data. These shocks first result in an increase and then a slow decline in future unemployment rates. Quarterly results that demonstrate this pattern have been extensively reported elsewhere for the U.S. and the U.K. in the post-war period.[1] It is not a pattern that results

[1] See Altonji and Ashenfelter (1980), and Ashenfelter and Card (1982).

Table 2

Persistence of a Unit Shock in Unemployment

Country	U.S.	U.K.	Canada	U.S.	U.K.	Canada
Period of Fit	1894 1983	1894 1983	1923 1983	1946 1983	1946 1983	1946 1983
Effect of the Shock in Period:						
0	1.000	1.000	1.000	1.000	1.000	1.000
1	1.169	1.190	1.185	.589	.982	.735
2	1.017	1.127	1.057	.183	.443	.46[illegible]
				.011		
3	.779	0.997	0.842	-.023	-.076	.293
4	.555	0.861	0.630	-.016	-.306	.182
5	.376	0.736	0.455	-.055	-.261	.113
6	.245	0.628	0.320	-.001	-.097	.070
7	.155	0.534	0.222	.001	.041	.043
8	.096	0.454	0.152	.000	.091	.027
9	.057	.386	0.103	.000	.068	.017
10	0.034	.328	0.069		.019	.010

Source: These are the moving average representations of the autoregressions in Table 1.

solely from movements in the labor force, because a similar pattern appears in quarterly U.S. employment data[2] and in monthly U.S. data on manhours worked.[3]

Second, the autoregressive (AR) structures for unemployment sometimes bear a striking similarity across countries when fit for the same time period. Although far from identical, the AR(2) representations for the U.S., the U.K. and Canada are quite similar over the period 1894-1983 in Table 1. As a consequence, the moving-average representations depicted in Table 2 are also similar.

Third, the exact empirical form of the persistence in the change in the unemployment rate does not appear to be temporally stable over long periods. This may be seen by comparing the AR(2) representations for the 1894-1945 period with the same representations for the 1946-83 period in Tables 1 and 2. The moving average representations have a typical humped shape in the longer period, but the impact of innovations to unemployment is more damped in the Post-War period.

B. Persistence in the Real Wage Rate

Table 3 reports selected estimates of autoregressions for the U.S. aggregate real wage rate. In the absence of any trend removal (columns 1 and 5) a simple random walk with drift (intercepts are not reported)

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[2]See Sargent (1978).

[3]See Kennan (1983).

Table 3

Autogressions for the Aggregate Real Wage Rate, U.S. Data

	1	2	3	4	5	6	7
Type of Data	Annual	Annual	Annual	Quarterly	Quarterly	Quarterly	Quarterly
Time Period	1929-76	1929-76	1929-76	1956-1980(1)	1956-1980(1)	1956-1980(1)	1956-1980(1)
Coefficient of:							
Real Wage Lagged Once	1.002	1.113	.953	1.100	.992	1.000	.868
(Standard Error)	(.007)	(.148)	(1.47)	(.110)	(.014)	---	(.068)
Real Wage Lagged Twice		-.111	-.196	-.22			
(Standard Error)		(.149)	(.140)	(.12)			
Lagged Change in Prices					-.297	-.341	-.425
(Standard Error)					(.134)	(.134)	(.139)
Linear Trend	No	No	Yes	Yes	No	Yes	Yes
Quadratic Trend	No	No	Yes	Yes	No	Yes	Yes
Seasonal Dummies				Yes	No	Yes	Yes

Source: Columns 1-3 are from Altonji and Ashenfelter (1980), Table 1. Column 4 is from Ashenfelter and Card (1982), Table V(a). The remaining columns are fit from Citibase data as described by Ashenfelter and Card (1982). The real wage is average hourly earnings in manufacturing divided by the consumer price index.

provides a very good fit to either quarterly or annual data. With trend removal, however, these autoregressions always imply damped, slowly decaying moving average representations.[4]

Columns 5-7 also report some very simple estimates of the effect of lagged price changes on the real wage rate. In the post-War quarterly data, at least, there is fairly strong evidence that rational forecasts of real wage rate changes may be a negative function of current price changes with an elasticity of -.3 to -.4. As we shall see, this is an important result, because it provides one simple potential link between a microeconomic model that emphasizes the role of real wage rates in the determination of labor supply and a macroeconomic model that admits a role for nominal price changes in the determination of labor supply.

C. Price Changes and Unemployment

Figures 1, 2, and 3 contain plots of the inflation/unemployment combinations for the years 1893-1945 for the U.S. and the U.K., and for the years 1921-1945 for Canada. Also indicated on each of these figures is the area within which the inflation/unemployment combinations for the years 1954-81 that are contained in Figures 4,5 and 6 would have fallen. Although it is natural to focus attention on the data for the later years contained in Figures 4,5 and 6, it is important to put these in context. First, it is clear that until recently, at least, the Post-War

[4]This result also holds up in monthly data. See Kennan (1983).

Figure 1

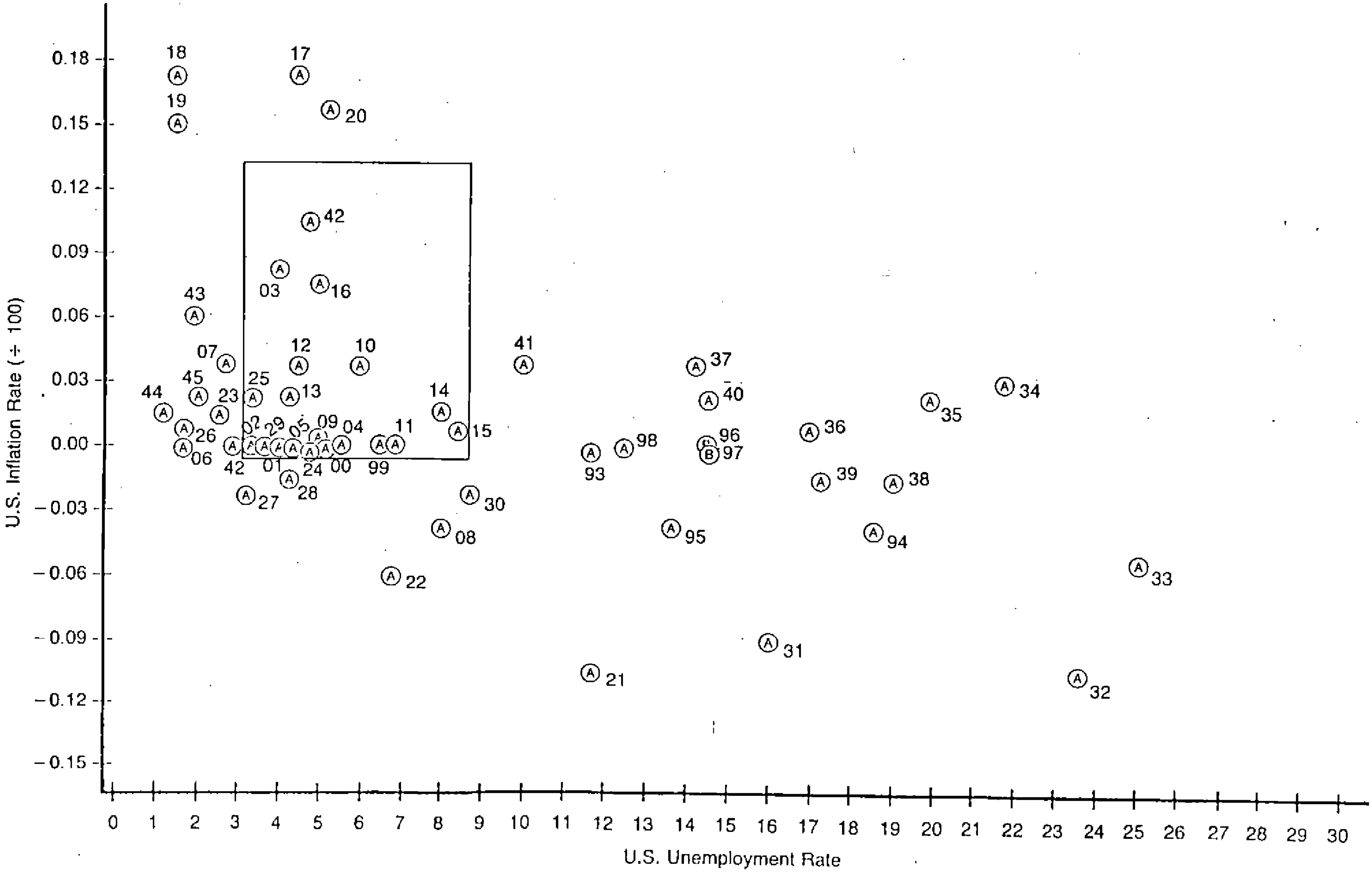
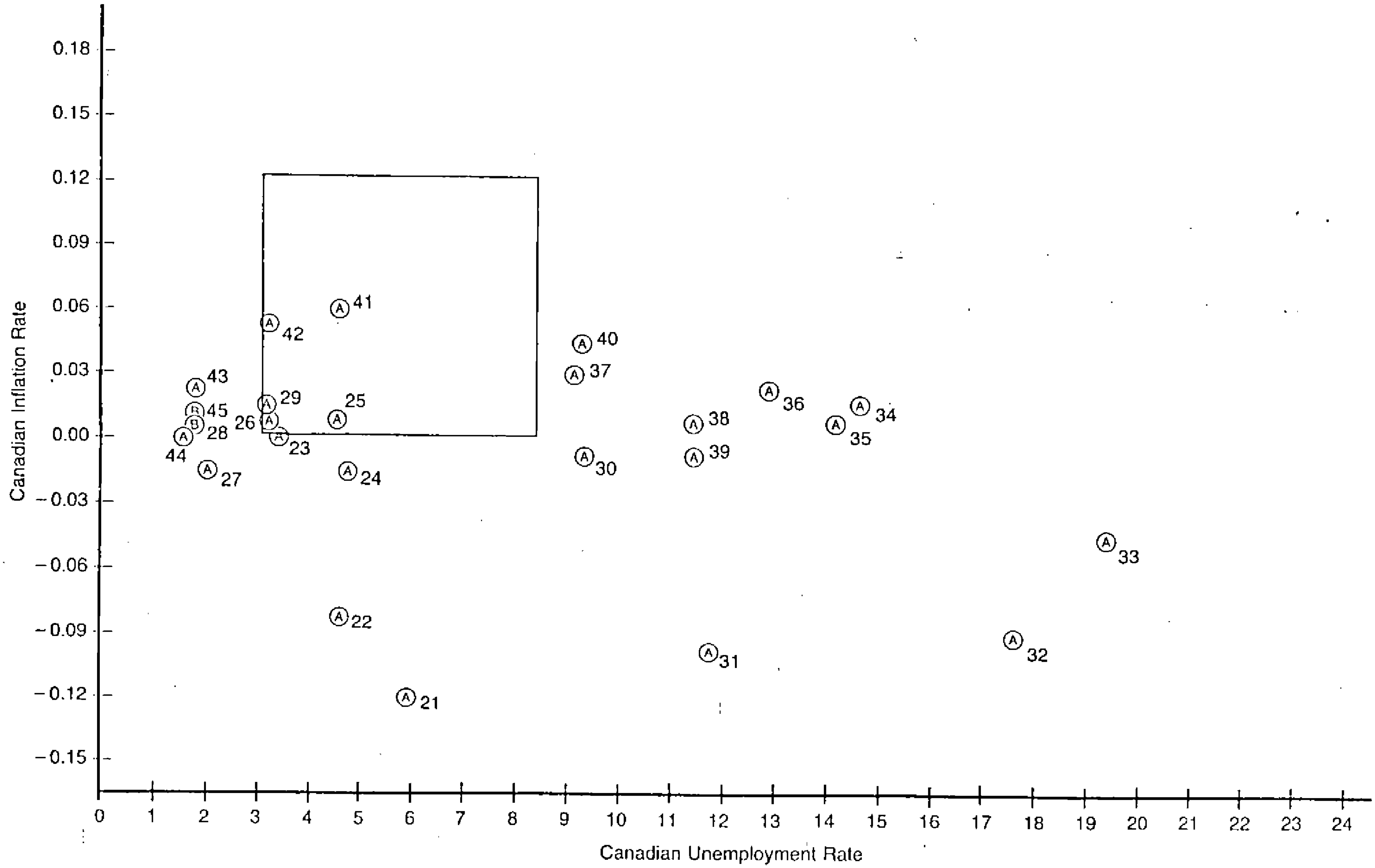

U.S. Inflation Rate (÷ 100)
0.18
0.15
0.12
0.09
0.06
0.03
0.00
−0.03
−0.06
−0.09
−0.12
−0.15
0 1 2 3 4 5 6 7 8 9 10 11 12 13 14 15 16 17 18 19 20 21 22 23 24 25 26 27 28 29 30
U.S. Unemployment Rate

Figure 2



Canadian Inflation Rate
0.18
0.15
0.12
0.09
0.06
0.03
0.00
-0.03
-0.06
-0.09
-0.12
-0.15
0 1 2 3 4 5 6 7 8 9 10 11 12 13 14 15 16 17 18 19 20 21 22 23 24
Canadian Unemployment Rate
A 41
A 42
A 40
A 37
A 43
A 29
25
A 36
A 34
B 45
26 A
A
A 38
B 28
A 35
A 44
A 23
A 39
A 30
A 27
A 24
A 33
A 22
A 32
A 31
A 21

Figure 3

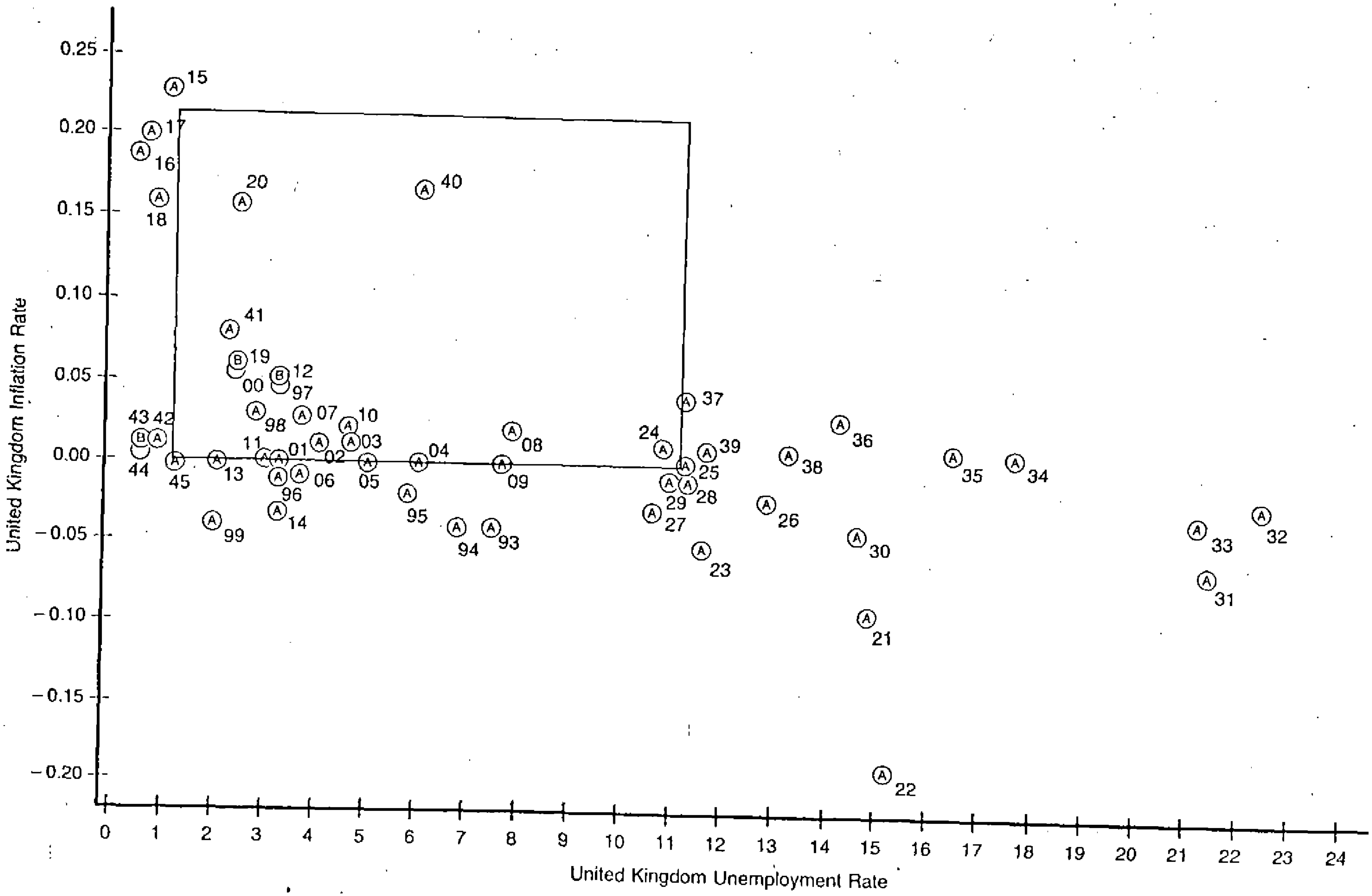

United Kingdom Inflation Rate
United Kingdom Unemployment Rate

Figure 4

American Data

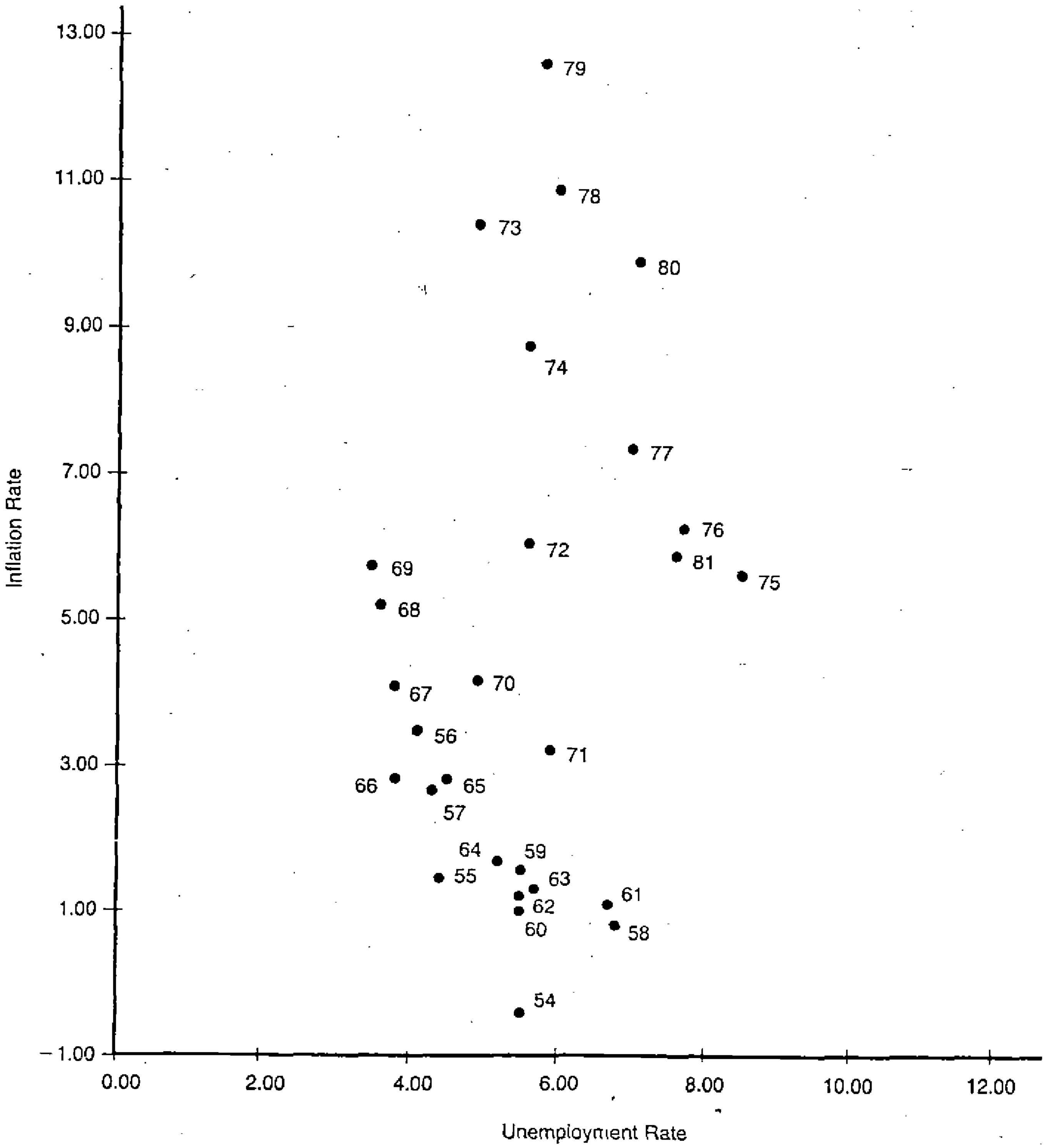


Figure 5
Canadian Data

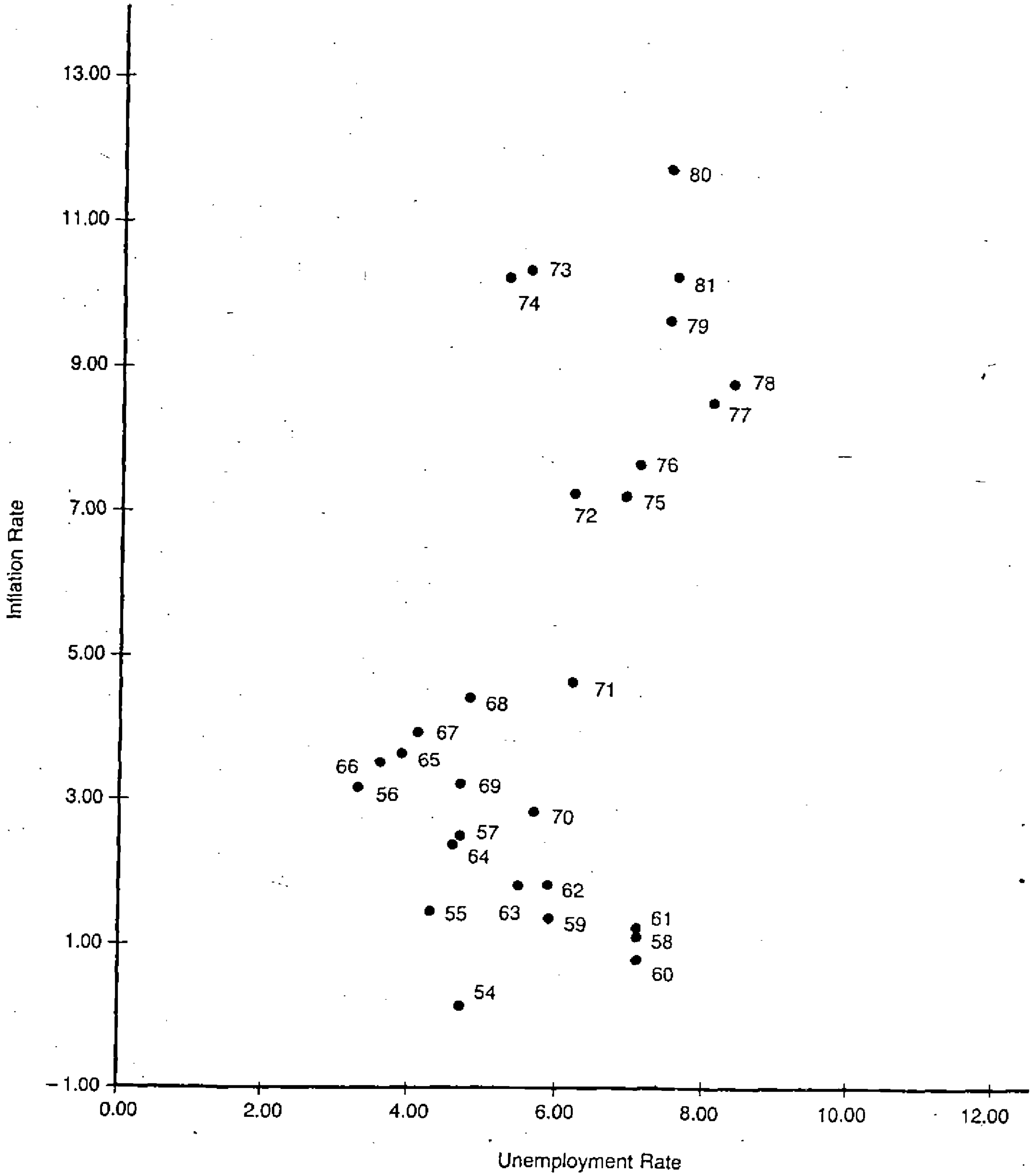


Figure 6

United Kingdom Data

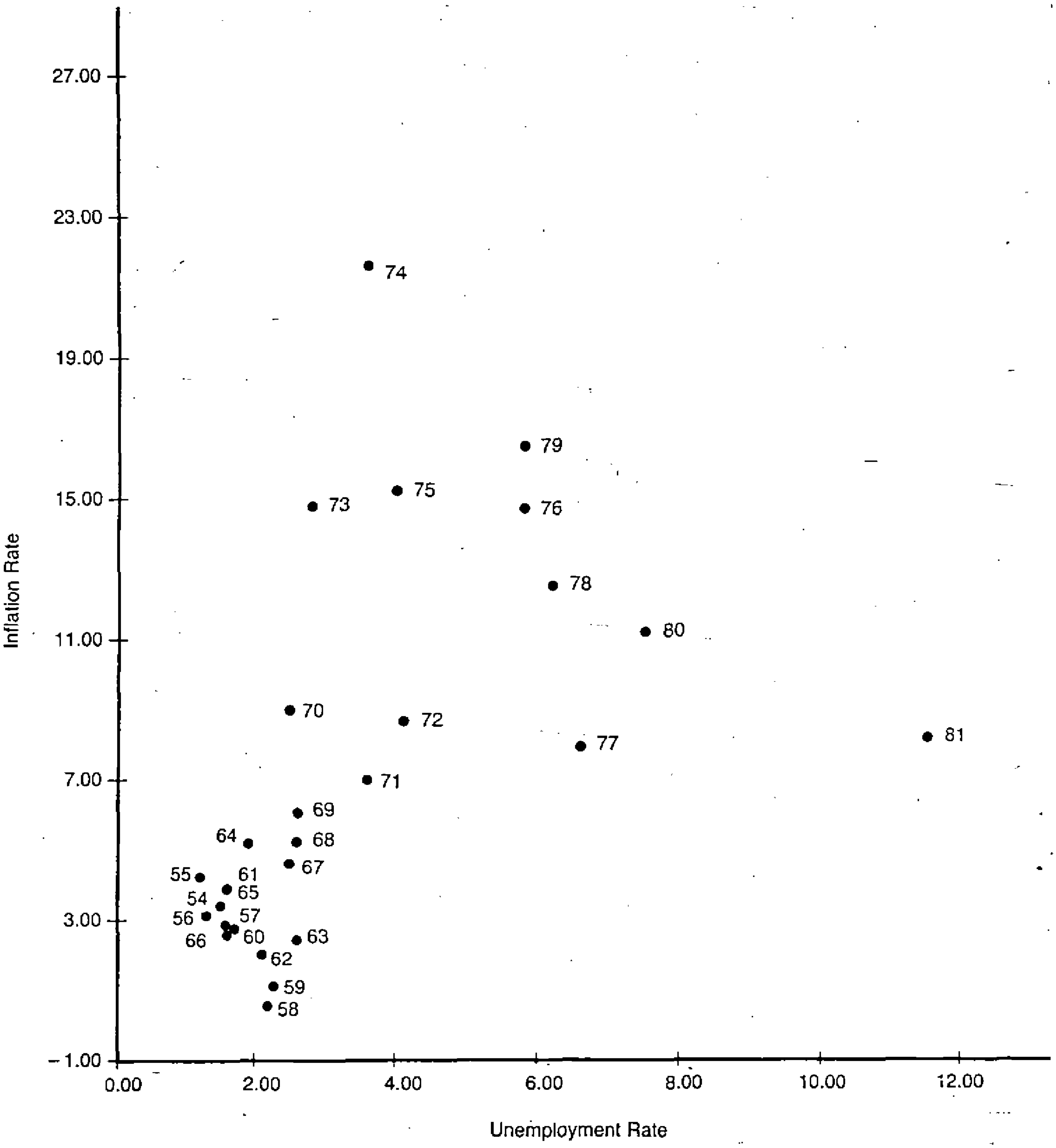
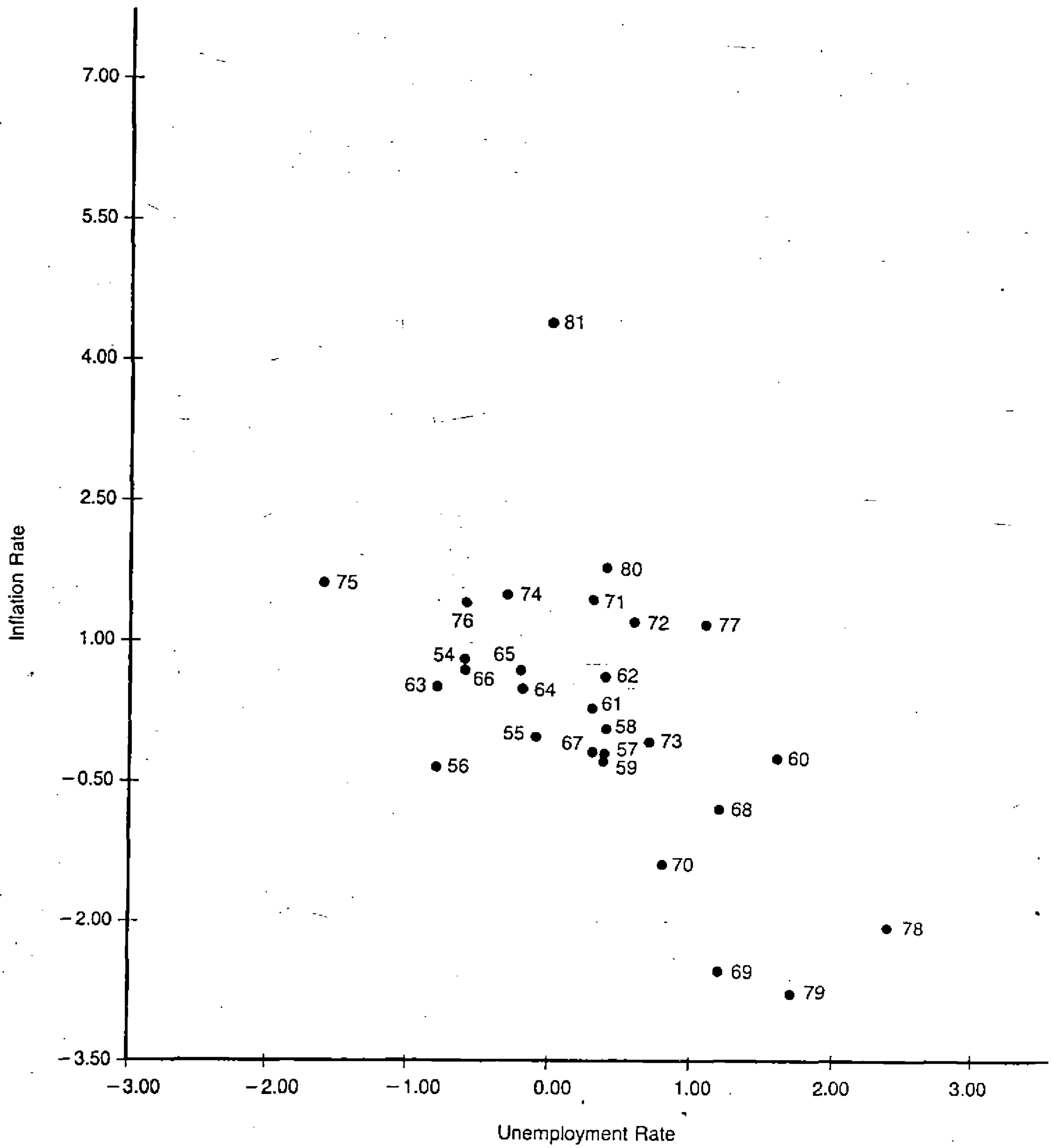


Figure 7

Differenced Data (Canada—U.S.)



range of U.S., Canadian, and U.K. unemployment experience has been historically small. Second, it is obvious that the relationship between inflation and unemployment appears very weak in the pre-War data. Examining these data alone would have been very poor preparation for the experience recorded in the 1950's and 1960's.

Indeed, it is especially interesting to consider how the contemporary inflation/unemployment data in Figures 4,5 and 6 would have looked to an observer in 1970. This is easy enough to do since all of the inflation/unemployment combinations after this year are in the upper right-hand half of the figures. It is extremely tempting to portray this as a sequence of short-run Phillips Curves that is continuously shifting up and to the right in these figures. Precisely how these short-run and long-run relationships can coexist is presumably one of the key questions that any aggregate model of the labor market is meant to address.

Of course, in the face of the extraordinarily weak contemporaneous correlations between inflation and unemployment that exist in both the pre-War and post-War data, a skeptic might simply deny the existence of any short run or long run relationship between these variables. One simple set of facts that I have found useful in confronting the skeptics is contained in Figure 7. This figure contains the scatter diagram of the difference between the Canadian and U.S. inflation rates against the difference between the U.S. and Canadian unemployment rates. The idea behind this comparison is a simple one, inspired, in part, by

Lucas's (1973) argument that price expectations might usefully be treated as a latent variable in the determination of the deviation of unemployment from its natural rate.[1] With similar slopes in their short run Phillips Curves and similar price expectations, apart from a random error, these inflation/unemployment differences will lie along a common short run Phillips Curve. The very loosely determined and negatively sloped empirical relationship in Figure 7 suggests that there may be something to this idea.

Table 4 contains the regression estimates of the relationship depicted in Figure 7, except that the coefficients on the inflation rate variable are unconstrained. As the table indicates, these coefficients are remarkably well determined and insignificantly different from each other. A similar analysis in Table 4 is somewhat less convincing when U.K. and U.S. employment rate differences are examined. The equality of regression coefficients on the U.K. and U.S. inflation rates is only striking when a quadratic trend is included in the regression equation.

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[5]For the i^{th} country write $u_{it} = \overline{u}_{it} + \alpha_i[\Delta\ln \overline{P}_{it} - \Delta\ln P_{it}]$, where u_{it} is unemployment, $\overline{u}_{it}$ is the natural rate of unemployment, $\ln P_{it}$ is the log of the price index, and $\ln \overline{P}_{it}$ is the log of the expected value of the price index based on information available at time period $t-1$. The crucial assumption is

$$\overline{u}_{it} - \overline{u}_{jt} + \alpha_i \Delta\ln \overline{P}_{it} - \alpha_j \Delta\ln \overline{P}_{jt} = y(t) + \varepsilon_t ,$$

with ε_t uncorrelated with $\Delta\ln P_{it}$ and $\Delta\ln P_{jt}$ and where $y(t)$ is a deterministic function that might be taken to represent trend-like shifts in differentials in natural rates or expected price levels. This condition would be satisfied, for example, with fixed anticipated exchange rates between Canada and the U.S., and Canadian prices determined largely in U.S. markets.

Table 4

Regressions of the Difference in Unemployment Rates on Inflation Rates[1]

Dependent Variable: Difference in Unemployment Rates in:	Quadratic Trend	Canadian Inflation Rate	U.S. Inflation Rate	U.K. Inflation Rate	R^2	Durbin-Watson Statistic
	Independent Variable:					
Canada and U.S	No	-.300 (.090)	.366 (.092)		.410	1.13
Canada and U.S.	Yes	-.441 (.129)	.365 (.092)		.472	1.40
U.K. and U.S.	No		.581 (.148)	-.171 (.091)	.485	1.19
U.K. and U.S	Yes		.236 (.113)	-.238 (.071)	.745	1.90

[1] Regressions are for 1956-1981.

In my view it is important to find simple structural methods for estimating the nature of the short run correlation between inflation and unemployment so that this relationship can be sorted out from shifts in the underlying determinants of structural (or "frictional" or "natural") unemployment. So long as this is not possible, the simple correlation between inflation and unemployment, whatever its sign, will continue to be used by the general public as an indication of causation. Just as the negative correlation between these variables in the 1950's and 1960's was taken to imply that an increase in inflation might reduce unemployment, the positive correlation that has materialized with the experience of the 1970's is now taken to imply that a reduction in inflation will reduce unemployment. I doubt whether either inference is the appropriate one to draw from these data, but in the absence of a simple and convincing demonstration to the contrary it is inevitable that more will be inferred from these correlations than is appropriate for wise public policy decisions. The simple analyses in Figure 7 and Table 4 are no doubt little more than an example of how a more convincing, but simplified analysis might proceed.

D. Aggregate Models of Labor Supply

As John Taylor (1983) observes in a recent survey, much of the last decade of research in macroeconomics has been inspired by a desire to produce explicit structural models that might rationalize the Phillips Curve observations in Figure 4. The line of research started by Lucas and Rapping (1969) and Lucas (1973) puts together a model of intertemporal labor supply with an assumption of rational expectations to do this job. Lucas and Rapping observe that demand shifts are the logical candidates for the cause of business cycle fluctuations in employment. Maintaining the assumption of continuous market clearing, however, requires that the short run labor supply curve be upward sloping or these demand shifts will result in real wage movements without corresponding movements in employment. Although long run labor supply is known to be generally insensitive to the real wage rate, or even backward bending [see Killingsworth (1983)], in a simple intertemporal model the short run elasticity of labor supply with respect to the real wage must be strictly positive.

To complete the Phillips Curve rationalization a correlation between short run movements in labor supply and changes in nominal prices must still be established. Either of two routes may be taken. In the case set out by Lucas (1973) workers may be incompletely informed about the aggregate nominal wage, the aggregate price level, or both. In this case workers may erroneously (but rationally) believe that their actual unexpected nominal wage increases (or price decreases) are "good draws" that will not be repeated. They will then want to capitalize on these good draws to a greater or lesser extent depending on the size of their intertemporal labor supply elasticity.

Alternatively, it may be the case that nominal prices or wages are a useful predictor of future real wage rates, for reasons otherwise unspecified. As Table 3 indicates, there is clearly some evidence based on the historical record that simple forecasts of future real wage rates might sensibly depend negatively on the current inflation rate, although the elasticity is small.[6] It follows that workers may reasonably assume that their current real wage rates are high, relative to what they may expect in the future, when inflation rates are high. Again, workers may use this information to capitalize on the "good draws" that higher current price inflation implies they may be getting if there is intertemporal substitution of labor supply.[7]

This intertemporal substitution model of the business cycle is internally consistent, complete, and it is clearly amenable to econometric testing. Most important of all, this macroeconomic model of the labor market is entirely consistent in its logic with the microeconomic models of labor supply behavior that have been the subject of considerable theoretical and empirical development over the last decade.[8] Indeed, research that establishes the empirical adequacy of the microeconomic models might even be taken to establish at least partial credibility for the macroeconomic model. So long as the microeconomists are satisfied with the empirical success of models that assume the con-

[6]This is also the case in Canada. See Card (1983).

[7]Detailed examples of this argument are contained in Altonji and Ashenfelter (1980) and Ashenfelter and Card (1982).

[8]Compare, for example, Altonji's (1982a) study of aggregate labor supply with his 1984 microeconomic study of life-cycle labor supply.

tinuous clearing of spot labor markets it appears that only one question remains: Are the microeconomic elasticities of labor supply big enough to explain the macroeconomic fluctuations?

As I have observed, the aggregate fluctuations in employment and unemployment are substantial. On the other hand, macroeconomic fluctuations in real wage surprises are apparently small. After all, the aggregate real wage is close to a random walk with a very small error variance. Alternatively, although price inflation rates are useful predictors of future real wage rates, this elasticity is also small. Apparently, intertemporal labor supply elasticities must be "large" if the macroeconomic model is to have any explanatory power.[9]

E. Empirical Tests of the Aggregate Model

With refreshing candor, even the earliest empirical tests of the intertemporal substitution explanation for aggregate fluctuations in employment, unemployment and inflation were not oversold. In an infrequently cited paper, Lucas and Rapping (1969, p. 349) describe their intertemporal substitution explanations for the Phillips Curve data by

[9]This ignores any role for real interest rates, but most of the post-War evidence suggests near constancy for real rates until the early 1970's. See Ashenfelter and Card (1982). Of course, this leaves open the role of real interest rates in explanations of employment and unemployment fluctuations since the early 1970's.

saying, "As reported econometric models go, ours can scarcely be called successful, but we think its failures are suggestive along several lines." They proceed to suggest that their findings are both empirically and theoretically consistent with a short run, but the absence of a long run, inflation/unemployment tradeoff. Their main conclusion, however, is that the estimated empirical relationships are highly unstable over time, and that a proper accounting of expectations may increase their explanatory power.[10]

This early empirical work by Lucas and Rapping predated the introduction of empirical methods for implementing the rational expectations hypothesis about expectation formation, and for a while their original empirical challenge was largely ignored. In an extremely thorough recent study, Altonji (1982) takes up the challenge and continues the careful empirical testing of the intertemporal substitution model in the rational expectations framework. Altonji's (1982, p. 784) main conclusion is that the empirical results for a long time-series of annual data in the U.S. do not support the intertemporal substitution model as a structural explanation for aggregate fluctuations in employment because, "For most specifications, the current real wage, the expected future real wage, and the expected real rate of interest are either insignificantly related to unemployment and labour supply or have the wrong sign." Although differing in detail, broadly similar conclusions on the empirical weakness of the intertemporal substitution model as an expla-

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[10]For the period 1946-65, for example, there is not a single regression coefficient on a wage or price variable that is much larger than its standard error in any regression that Lucas and Rapping report.

nation of the aggregate data were reached early on by Sargent (1973), and subsequently by Altonji and Ashenfelter (1980), Andrews and Nickell (1982), Ashenfelter and Card (1982), and Kennan (1983). In all of these studies, a major difficulty is the identification from the data of an intertemporal labor supply elasticity that is (positive and) large enough to reconcile the dramatically greater fluctuations in employment than in wage rates that are observed.

There are, of course, many acknowledged difficulties in extracting labor supply elasticities from aggregate time-series data. These include problems of aggregation, simultaneity in the determination of employment and wage rates, structural shifts in monetary policy rules and in government employment policies, and the like. No doubt partly to overcome these difficulties, it is natural to turn to microeconomic studies of labor supply to see whether they have met with greater empirical success and have resulted in more precise estimates of key parameters. If some important parameters can be determined from the microeconomic studies, these may naturally be included as the building blocks in more persuasive aggregate models of employment fluctuations.[11] The question then remains as to how well these microeconomic models of

[11] See Kydland and Prescott (1982) for an example of the application of such methods. In many ways this is also the spirit of Kennan's (1983) continuing work, although his approach is to "estimate" supply and demand parameters from the aggregate data and then inspect them for their reasonableness.

intertemporal labor supply have survived empirical testing and estimation.

II. Microeconomic Evidence of Intertemporal Substitution in Labor Supply

The empirical setting for microeconomic analyses of intertemporal movements in labor supply must clearly be the life cycle. In fact, setting out the simple theoretical models available serves two purposes. First, it shows the clear connection between the life-cycle labor supply model and the permanent income theory of consumption. Indeed, the latter is simply the consumption plan derived from the former, and permanent income is nothing more than the appropriately discounted present value of future wage rates. Second, to be tractable, the empirical analyses are going to require some form of linearity and some simple method for summarizing a consumer-worker's future prospects. Intertemporally additive utility and special functional forms are necessary to provide the justification for these simplifications.

A. Life-Cycle Labor Supply with Perfect Foresight

One simple model that generates a linear earnings function is the Stone-Geary utility function. In an intertemporal context this function is additive both at a point in time and over time. In particular, consider the utility function

$$v = \sum_t (1 + \rho)^{-t} \, [B_1 \ln(\gamma_h - h_t) + B_2 \ln(c_t - \gamma_c)] \, , \tag{1}$$

where ρ , B_1 , B_2 , γ_h , and γ_c are parameters and h_t and c_t are hours of work and aggregate commodity consumption. In this setup there is a minimum necessary commodity consumption level γ_c , a maximum feasible hours of work level γ_h , and a rate of time preference ρ , all of which are constant over time.

Maximizing (1) subject to the lifetime budget constraint

$$\sum_t (1 + r)^{-t}(w_t h_t + y_t - p_t c_t) = 0$$

with fixed interest rate r , unearned income y_t , and assuming for simplicity that the rate of time preference $\rho = r$, leads to the first order conditions

$$B_1/(\gamma_h - h_t) = \lambda w_t$$

$$B_2/(c_t - \gamma_c) = \lambda p_t \text{ .}$$[12]

Using the normalization $\sum_t(1 + r)^{-t}(B_1 + B_2) = 1$, these lead to the explicit solution for λ of

$$\lambda = [\Sigma(1 + r)^{-t}(y_t + \gamma_h w_t - \gamma_c p_t)]^{-1} ,$$

the labor earnings functions

$$(2) \qquad w_t h_t = \gamma_h w_t - B_1 \lambda^{-1} ,$$

[12]The case $\rho \neq r$ is worked through by Ashenfelter and Ham (1979). It leads to the multiplication of λ^{-1} in (2) and (3) by the term $(1 + r)/(1 + \rho)^t$ and a re-normalization of the income derivatives.

and the consumption functions

$$(3) \qquad p_t c_t = \gamma_c p_t + B_2 \lambda^{-1} .$$

In this setup real consumption will remain constant if the real price of consumption is unchanged, and will be proportional to the real discounted present value of discretionary income, λ^{-1} , if minimum consumption requirements, γ_c , are negligible.

Intertemporal movements in labor earnings in (2) are solely a result of life cycle or time-series movements in w_t . These movements are governed entirely by the parameter γ_h . To see this, note that λ^{-1} is a constant in (2) so that changes in labor earnings are

$$(4) \qquad \Delta w_t h_t = \gamma_h \, \Delta w_t .$$

Indeed, the proportional change in earnings is

$$\Delta(w_t h_t)/w_t h_t = (\gamma_h/h_t)[\Delta w_t/w_t] .$$

In this model the so-called intertemporal elasticity of labor supply is therefore $(\gamma_h/h_t) - 1$. Since $\gamma_h > h_t$ is required for convexity of the worker's indifference curves, this implies that the intertemporal elasticity of labor supply must be non-negative. This result does not constrain the income effects in this model at all. It is in this sense that the intertemporal additivity of the model does not constrain the data to produce "small" intertemporal elasticities of labor supply.[13]

Ham and I have fit this model to data from the Panel Survey of Income Dynamics (PSID), but before reporting those results let me turn to some data that provide a very simple method for estimating γ_h. This scheme is based on the observation that γ_h in (4) is essentially a regression without a constant term. One consistent estimator for γ_h is therefore the ratio of the mean of $\Delta(w_t h_t)$ to the mean of Δw_t. The advantage of this estimator is that it remains consistent even when zero-mean measurement errors are appended to w_t and $w_t h_t$ in equation (2).[14]

To get a feeling for the estimates obtained in this way consider the mean changes in real earnings and real wage rates reported from the PSID in Table 5. The third column reports the estimates of γ_h, while the fourth column reports the average of mean hours worked in the two years considered. There are two disturbing features about these estimates of

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[13] Indeed, the effect of a lifetime change of Δw_t on labor earnings is

$$\Delta w_t h_t = \gamma_h[\Delta w_t - B_1 \sum (1+r)^{-s} \Delta w_s],$$

so that the long run labor supply elasticity is $(\gamma_h/h_t)\,[1-B_1 \sum (1+r)^{-s}]-1$. This long run labor supply elasticity must be smaller than the intertemporal labor supply elasticity, $\gamma_h/h_t -1$, and its size will depend on the income derivative and the length of the time horizon; it may, of course, be zero.

[14] In fact, this is precisely Wald's (1940) method for "fitting a straight line if both variables are subject to error." Consistency of this estimator demands primarily that the probability limit of the denominator in this ratio be non-zero.

Table 5

Changes in Real Earnings and Real Wage Rates,

Panel Survey of Income Dynamics

(White Males, 25-50 years old in 1967)

	Change in Real Earnings	Change in Real Wage	γ_h	Mean Hours Worked
1967-68	486	.19	2,558	2,416
1968-69	313	.20	1,565	2,403
1969-70	-101	.01	-10,100	2,370
1970-71	206	.18	1,144	2,352
1971-72	561	.12	4,675	2,367
1972-73	396	.16	2,475	2,370
1973-74	-371	-.02	18,550	2,328

Source: Appendix of Ashenfelter and Ham (1979); earnings and wage rate deflated by consumer price index, 1967 = 1.0.

γ_h. First, they are very unstable. Second, in three of the years considered they are lower than the actual mean of hours worked. Discarding as extreme outliers the results for 1969-70 and 1973-74 leads to an average ratio of the ratios $\gamma_h/h_t = 1.04$, which implies an intertemporal labor supply elasticity of .04. Obviously, with these data virtually any estimate of γ_h may be obtained depending on what the empirical analyst wants to see.

There are clearly difficulties in using this simple estimator to calibrate the size and stability of the intertemporal labor supply elasticity. Perhaps most disturbing are the possibility that aggregate supply shocks or their determinants will obscure movements along the supply schedule (4). More generally, anything that might successfully and correctly be removed from the panel data by the addition of year dummy variables will produce a specification bias in these results. It is important to emphasize, however, that the consistency of many of the estimates of the intertemporal labor supply elasticity that I report below are dependent on the same assumptions necessary to ensure the consistency of the estimates of γ_h in Table 5 <u>and</u> typically on further assumptions.

Ham and I have also fitted equation (4) to the PSID micro data directly. These results are even more disappointing. We obtained very precise estimates of γ_h based on the pooled covariances in the data of around 1,900 hours. This result also implies a negative intertemporal elasticity of labor supply.

A different model that leads to a log linear labor supply function has been suggested by Heckman and MaCurdy (1980) and MaCurdy (1981). As Abowd and Card (1983) observe, with a constant real interest rate and negligible initial assets this model leads to precisely the log linear labor supply function initially proposed by Lucas and Rapping (1970). Taking

$$v = \sum_t (1 + \rho)^{-t} [\ln c_t - g(h_t)] ,$$

where

$$g(h_t) = \exp\{-A(1 + \eta)/\eta\}[\eta/1 + \eta] h_t^{(1 + \eta)/\eta}$$

leads to the labor supply functions

$$\begin{aligned} (5) \qquad \ln h_t &= A(1 + \eta) + \eta \ln w_t + \eta \ln \lambda \\ &= A + \eta[\ln w_t - (1 - \beta)\sum_s \beta^s \ln w_s] , \end{aligned}$$

where $\beta = 1/(1 + r)$.

In this setup η is the intertemporal labor supply elasticity and must be non-negative. As before, the constancy of $\ln \lambda$ implies that the proportionate change in labor supply over the life cycle is governed by

$$(6) \qquad \Delta \ln h_t = \eta \Delta \ln w_t .$$

Again, there are some straightforward estimates of η available from the ratios of the means of $\Delta \ln h_t$ to $\Delta \ln w_t$ in a panel of data. To provide a feeling for the size of these estimates I report in Table 6

Table 6

Changes in Log Real Earnings, Log Hours, and Log Real Wages,

Panel Survey of Income Dynamics

(Males Heads of Households, 21-64)

Date	Change in Log Earnings	Change in Log Hours	Change in Log Wage	η[a]	Change in Unemployment Proportion[b]	Change in Log Manhours[c]
1969-1970	.032	-.011	.043	-.26	.009	-.02
1970-1971	.030	.003	.027	.11	.007	-.01
1971-1972	.072	.021	.051	.41	-.004	.04
1972-1973	.048	.021	.027	1.29	-.007	.04
1973-1974	-.051	-.042	-.009	4.67	.006	.00
1974-1975	-.041	-.027	-.014	1.93	.023	-.05
1975-1976	.046	.012	.034	.35	-.008	.04
1976-1977	.024	.002	.022	.09	-.006	.04
1977-1978	.005	-.003	.008	-.38	-.007	.05
1978-1979	-.055	-.042	-.013	3.23	.001	.03

Source: Abowd and Card (1983), Table 2

Employment and Training Report of the President, Tables A-30, C-13, 1982.

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a/ Calculated as the ratio of the mean change in log hours to the mean change in log wages.

b/ Change in unemployment proportion for males aged 35-44.

c/ Change in the logarithm of the payroll series data on the aggregate weekly hours index.

the data on the change in log hours and log wages from the PSID computed by Abowd and Card (1983).

The estimates of η in Table 6 are qualitatively consistent with the results in Table 5 except for the year 1970-71. In general, however, the data in Table 6 are far more congenial to an estimate of the intertemporal labor supply elasticity that is positive and large in magnitude. Only two of the ten estimates of η are negative, and the simple average of all the estimates is 1.14. Deleting the two extreme outliers leads to an estimate of the intertemporal labor supply elasticity of .89. (This is equivalent to deleting the two estimates with the denominators closest to zero in absolute value.) As before, however, these estimates are very unstable and this instability casts serious doubt on the credibility of this model.

There are several ways to use the covariances in the data to estimate η. The simplest method is simply to compute the regression coefficient of $\Delta \ln h$ on $\Delta \ln w$. Abowd and Card (1983) report all of the necessary data to do this from the PSID and from the National Longitudinal Survey of Older Men (NLS). Although it does not appear to be widely reported in the literature, this regression coefficient is always negative <u>and</u> significantly different from zero at quite small probability levels. In the PSID it is $-.36$ and in the NLS it is $-.28$

for the data reported by Abowd and Card (1983).[15]

I do not want to suggest that it is impossible to use the covariances in the data to find a regression coefficient with the sign implied by equation (5). For example, MaCurdy (1981) observes that adding $\eta\Delta \ln h_t$ to both sides of (6) produces the relationship

$$(7) \qquad \Delta \ln h_t = (\eta/1 + \eta)\Delta \ln w_t h_t \quad .$$

This suggests computing the regression coefficient of the change in the log of hours on the change in the log of earnings. In the PSID and NLS data these regression coefficients imply estimates of η of around .78 and .61 , respectively.

It does seem clear, however, that simple applications of either the linear earnings equation (4) or the log linear hours equation (6) will require some subtle manipulation of the data before they will produce credible estimates of the intertemporal labor supply elasticity. As a result, these estimates are likely to be sensitive to the model specified, although preliminary indications are that they are not likely to be larger than .7 or .8.

B. Models of Life-Cycle Labor Supply with Measurement Error

The presence of measurement error has been suggested as one important reason for modifying equations (4) and (6). One suggestion is to

[15] I first became aware of this "fact" after seeing Altonji's (1984) two estimates of this regression coefficient. Using two different measures of the wage rate from the PSID, Altonji reports estimates of the regression of $\Delta \ln h_t$ on $\Delta \ln h_t$ of −.40 and zero. The former is essentially based on the same data as reported by Abowd and Card (1983).

recognize the presence of measurement error in both $\Delta \ln w_t$ and $\Delta \ln h_t$ at the micro level. As I have observed, the simple ratio of means estimates in Tables 5 and 6 need not suffer from bias induced by measurement error. On the other hand, the usefulness of this simple procedure depends critically on the assumption that unmeasured economy wide shocks to real interest rates or other aggregate variables can be safely ignored. Using the covariances in the panel data with time means subtracted out does not run this risk.

MaCurdy (1981) circumvents these issues by estimating equation (5) by an instrumental variables scheme. With time means subtracted out of the data his estimates of the regression coefficient of $\Delta \ln h_t$ on $\Delta \ln w_t$ are .10 and .15 with standard errors of about the same magnitude. His estimates of the regression coefficient of $\Delta \ln h_t$ on $\Delta \ln w_t h_t$ imply estimates of η of .45 and .30 with standard errors of about two-thirds these magnitudes. These are not large elasticities and the imprecision of their estimation is disturbing. The imprecision no doubt results from the inevitably poor quality of what are essentially time-invariant instrumental variables.[16]

In an extremely thorough empirical study Altonji (1982b) reports several efforts to account for measurement error in an attempt to estimate (6). He reports three alternative sets of results from the PSID data. The first set uses an instrumental variables scheme designed to reproduce MaCurdy's results. The estimated intertemporal labor

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[16]It is well known that wage rates in a cross-section are roughly a semilogarithmic function of schooling, experience, and experience squared. The first-difference in the log wage is therefore approximately a linear function of experience, the main instrument available.

supply elasticity falls in the range .08 to .50 depending on whether time means are subtracted out of the data and whether age is included in the labor supply equation. Estimated sampling errors fall in the range .12 to .4, however, so that elasticities are still imprecisely estimated. A second procedure uses an alternative (but contemporaneously measured) wage variable as an instrument for the wage in a classical instrumental variables set up for handling measurement error. With this procedure the estimates of the intertemporal substitution elasticity are around .04 to .07, depending on specification. Estimated sampling errors are very small also, at around .07, so that substitution elasticities larger than .25 may be ruled out. In a third procedure Altonji recognizes that a contemporaneously measured alternative wage variable may be contaminated by common measurement errors or, in a model with uncertainty, correlated with labor supply function errors. Using a lagged alternative wage variable Altonji estimates intertemporal labor supply elasticities around .05, but estimated sampling errors are now around .45. All of these estimates rule out substitution elasticities greater than unity. Altonji concludes that these estimates suggest an intertemporal labor supply elasticity in the range 0 to .35, although I prefer to state all these results and their limitations so that they speak for themselves.

Altonji (1984) also presents estimates of η based on a procedure that exploits the marginal condition for the consumption plan associated with equation (5). The basic idea is that the marginal condition for optimal consumption requires that $1/c_t = \lambda p_t$, so that equation (5) can also be written as

$$\text{(5a)} \qquad \ln h_t = A + \eta \ln w_t - \eta \ln p_t c_t \ .$$

Differencing this equation and using data on food consumption in the PSID, and an instrumental variables procedure, Altonji estimates η in the range of .04 to .30 under various specifications. These estimates have somewhat smaller sampling errors than those reported previously, so that large estimates of the intertemporal labor supply elasticity may be ruled out if these estimates are accepted.

Abowd and Card (1983) have presented some persuasive evidence that much of the variation in both hours and earnings in the available longitudinal data may be a result of measurement error. To see the nature of this evidence consider any measured variable z_t^* whose true value is z_t . Suppose that measurement error e_t is serially uncorrelated, that e_t is uncorrelated with z_t , and that Δz_t is serially uncorrelated. Then $\text{cov}(\Delta z_t^*, \Delta z_{t-1}^*,) = -\sigma_e^2$ and $\text{var}(\Delta z_t^*) = 2\sigma_e^2$ so that the first-order autocorrelation coefficient of Δz_t^* is $\text{cov}(\Delta z_t^*, \Delta z_{t-1}^*)/\text{var}(\Delta z_t^*) = -1/2$, and the serial correlation coefficients at all higher order lags are zero. In effect, these assumptions imply that Δz^* is a first order moving average process of the form

$$\text{(8)} \qquad \Delta z_t^* = \varepsilon_t - \varepsilon_{t-1}$$

Abowd and Card present data that imply first-order autocorrelation coefficients for hours and earnings in the PSID data of -.35 and -.34.

Neither the second nor third order autocorrelation coefficients in their data are as large (in absolute value) as -.04. Although this is hardly conclusive, it suggests the possibility that a substantial fraction of the panel data movement in hours and earnings may be composed of measurement error. Indeed, Abowd and Card take the null hypothesis against which they test equation (6) to be a simple model of measurement error much like equation (8). Although they reject this model as a complete explanation for the data, a major message in their paper is the importance of dealing with this problem.

Altonji (1984) provides further evidence of measurement error in the main wage series used in a typical panel data study of labor supply. The change in the conventional wage measure in these studies is Δw^*_t, the change in the ratio of labor earnings to annual hours at work. For hourly workers in the PSID the change in the hourly wage rate, Δw^{**}_t, is also recorded. Assuming that both Δw^*_t and Δw^{**}_t are additive combinations of the change in the error free wage, Δw_t, and independent measurement errors, e^*_t and e^{**}_t, then

$$\Delta w^*_t = \Delta w_t + e^*$$

$$\Delta w^{**}_t = \Delta w_t + e^{**}_t$$

It follows that $\text{cov}(\Delta w^*_t , \Delta w^{**}_t)/\text{var}(\Delta w^*_t)$ and $\text{cov}(\Delta w^*_t \;\; \Delta w^{**}_t)/\text{var}(\Delta w^{**}_t)$ give measures of the fractional components of the two wage change measures that are not mesurement error. In Altonji's PSID data $\text{cov}(\Delta w^*_t , \Delta w^{**}_t) = .0049$, $\text{var}(\Delta w^*_t) = .0498$, and $\text{var}(\Delta w^{**}_t) = .0177$ This implies that 90 percent of the variance in Δw^* and 72 percent t

of the variance in Δw_t^{**} is measurement error. Moreover, since var (Δw_t^{**}) > var (Δw_t), it follows that $1-\text{var}(\Delta w_t^{**})/\text{var}(\Delta w_t^{*})$ is a lower bound on the percentage of the variance in Δw_t^{*} due to measurement error of around 60 percent. I do not mean to suggest that the wage change data contain only mesurement error. Still, it seems clear that the longitudinal data series available for identification of an intertemporal labor supply elasticity are very noisy.

C. Experimental Evidence on Labor Supply

By now there exist several experimental studies of labor supply, at least one of which is designed to address the importance of the life-cycle model as an explanation for hours changes. These studies are experimental in the sense that families are randomly assigned to a negative income tax treatment or a control group. The Seattle-Denver program is by far the largest, and it was designed explicitly to address the effect of negative income tax programs of different lengths on labor supply. It is precisely the possibility that a transitory negative income tax program might have different effects on labor supply from a permanent program that is at the heart of the life-cycle model. From the standpoint of judging the transfer costs of a negative income tax program, of course, this issue is mainly of concern in determining the extent to which a short term experiment simulates the impact of a long term program, and this was the public policy issue being addressed.

Some very straightforward estimates of treatment effects from this experiment are contained in Table 7. These are computed from coefficients on dummy variables that describe the particular treatments indicated and that contain only pre-experimental measures of other variables in the regressions. All of these estimates indicate a decline in hours at work among the treatment groups relative to the control groups, and many of these declines are statistically significant.

Since these are short duration programs, to the extent that the life-cycle is the basis for decision-making the tax effects should be exaggerated compared with what would be observed as the result of a permanent wage change. For example, in the Stone-Geary model the labor supply elasticity with respect to a permanent wage change (ignoring discounting) is $(\gamma_h/h_t)(1-B_1 n)-1$, where n is the length of the life-cycle. This long run labor supply elasticity is strictly smaller than the intertemporal labor supply elasticity, γ_h/h_t-1.

The data in Table 7 are not easily explained by the life-cycle model, however. First, the five year program appears to have a considerably greater effect on labor supply than does the 3 year program. This suggests that income effects must be an important component of the labor supply response. On the other hand, labor supply effects after completion of both programs are essentially negligible, which suggests that these income effects have not been distributed smoothly over the life-cycle in accord with the predictions of the model.

Another disturbing feature of the results in panel B of Table 7 concerns the apparent tax effects in the data. This panel shows three two

Table 7

Percentage Effects (Relative to Control Group) of Various Negative Income Tax Plans on Husbands Hours of Work, Seattle-Denver Income Maintenance Experiment

A. Effects by Duration of Program

	Years After Start of Program					
	1	2	3	4	5	6
Duration of Program:						
3 years	-1.6	-7.3**	-7.3**	-.5	-.2	
5 years	-5.9**	-12.2**	-13.2**	-13.6**	-12.3**	+3.0

B. Effects by Guarantee and Tax Rate for 2nd Program Year

Guarantee Level:	Tax Rate:	
	.50	.70
$3,800	-6.7	-5.6
4,000	-8.8**	-1.5
5,600	-11.8**	-10.4**

Sources: Final Report of the Seattle-Denver Income Maintenance Experiment, Vol. 1, SRI International, May 1983, Tables 3.4 and 3.9; and Overview of the Seattle-Denver Income Maintenance Experiment Final Report, Office of Income Security Policy, U.S. Department of Health and Human Services, May 1983, Table 4.

*Indicates significantly different from zero at the .05 level.

**Indicates significantly different from zero at the .01 level.

way contrasts (at different guarantee levels, which is the income received at zero work hours) of 70 versus 50 percent tax rate programs. In every case the decline in work effort is smaller under the higher tax rate program. This also suggests that income effects are playing a far more important role in these data than would be consistent with a major role for intertemporal substitution.

Surprisingly, research to date has only scratched the surface of what is possible with the experimental data available. A careful analysis of these data for the purpose of exploring and testing the detailed predictions of the life-cycle model of labor supply is long overdue.

III. Conclusion

In my view the microeconomic empirical work based on the life-cycle model of labor supply represents one of the finest amalgams of careful data analysis and applied theoretical work that exists in modern economics. It seems clear, however, that these analyses of intertemporal labor supply have not yet produced a coherent empirical explanation of the aggregate movements in hours of work in the available panel data. A simple way to see this is to first examine the time-series movements in the average hours at work in column 2 of Table 6 (for the PSID). The basic idea of the intertemporal labor supply model is to explain these changes as movements along a fixed supply curve. There are three problems in doing this. First, average hours and the average real wage must move in the same direction, which, as the table indicates, occurs most, but not all of the time. Second, as column 4 indicates, the slope of the labor supply function must fluctuate considerably from year to year in order to square up the aggregate hours and wage rate changes. Finally, the average labor supply elasticity must apparently be quite large to square up these hours and wage rate movements, while the available estimates of its slope that I have surveyed are, in fact, very small. The basic empirical problem seems to be that within the life-cycle, the person-specific correlation between hours and wages is simply too small to explain the time-series movements in average hours relative to the time-series movements in average wage rates. The intertemporal substitution hypothesis originally advanced by Lucas and Rapping was, of course, precisely the suspicion that this was not the case.

It remains to consider what light these findings shed on the more familiar data that register movements in the business cycle and to which macroeconomists are accustomed. Apparently the connection is not very straightforward, as the last two columns of Table 6 indicate. The fifth column of this table contains the annual first difference of the male unemployment proportion for workers aged 35-44. This is an often used cyclical indicator and it moves in tune with the unemployment rates for most other groups. A comparison of the PSID hours changes in column 2 of the table with these unemployment changes (they are in a similar scale) indicates that these series are by no means identical. To be sure, both series indicate similar magnitudes for the 1974-75 recession, but they move very differently in the three preceding years and in the last year of the sample. The last column of Table 6 shows that a comparison of the PSID average hours data with an index of aggregate manhours from the BLS payroll data fares no better. These latter data are no doubt heavily influenced by demographic and other trends, but it is by no means clear that any simple detrending will reconcile them with the PSID data. It appears that a careful reconciliation of the basic microeconomic and macroeconomic data is going to be necessary before further conclusions are warranted. Such a research project deserves high priority for future research.

In tracing through the status of the empirical research on microeconomic models of life-cycle (or intertemporal) labor supply, it is difficult to come away with the impression that these simple implementable models are providing good descriptions of the available longitudinal data.[17] In this regard it seems that the macroeconomic and microeconomic models of labor supply share much in common. Of course, it is always possible to attribute the empirical difficulties of both types of models to measurement error or flaws in functional forms. It is even possible that the measurement error in the microeconomic data is so severe that there is little or nothing to be gained from any analysis of it. Likewise, it may be that the restriction to linear functional forms is too restrictive, although it is hard to imagine

[17]Others have come (independently, I should add) to similar conclusions. Pencavel (1984, p. 147) writes, after a survey of the intertemporal labor supply research: "...the greater part of the variations in male labor supply across workers and over time is left unexplained by this research. A great deal of effort has been brought to bear on what appear to be relationships of second-order of importance."

that the kind of data currently available could credibly support anything more elaborate.

It is important to emphasize that the empirical difficulties in estimating intertemporal labor supply elasticities are not a problem primarily at the macroeconomic level. If there is something missing from these models of labor supply it is apparently missing both from the microeconomic and the macroeconomic models. What might it be?

In my view it must surely be the case that the long run behavior of average hours at work are mainly a result of worker preferences as between the consumption of goods and leisure. At the same time, it seems reasonable to suppose that demand induced movements in hours worked, at predetermined wage rates, may likewise have a role to play in the short run interplay of hours and earnings determination.[18] It should be emphasized that the existence of such demand related hours shocks has no particular normative implication for any public policies. Moreover, demand shocks as an explanation for employment fluctuations date at least to Adam Smith, who wrote:

> . . . the wages of labour in different occupations vary with the constancy or inconstancy of employment. Employment is much more constant in some trades than in others. In the greater part of manufactures, a journeyman may be pretty sure of employment

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[18]The empirical implications of the supply side of such a model have been worked out in Abowd and Ashenfelter (1979, 1981), but the empirical models estimated in those papers are not notably successful.

> almost every day in the year that he is able to work. A mason or bricklayer, on the contrary, can work neither in hard frost nor in foul weather, and his employment at all other times depends on the occasional calls of his customers. He is liable, in consequence, to be without any. What he earns, therefore, while he is employed, must not only maintain him while he is idle, but make him some compensation for those anxious and desponding moments which the thought of so precarious a situation must sometimes occasion.

It seems clear that Smith did not expect the wage rates of masons to fluctuate with hard frost or the occasional calls of their customers, and that he did not expect to explain employment fluctuations as a response to such wage rate fluctuations. Constructing a testable model that might reconcile the long run determination of hours worked by worker preferences with the short run interaction of observed employment and earnings may be the missing ingredient in both the macroeconomic and microeconomic models of labor supply.

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