THE EMPIRICAL SENSITIVITY OF THE PHILLIPS CURVE

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Few economic concepts have become as embedded in the calculus of governmental decision-making as has the notion of a trade-off between inflation and the full employment of (labour) resources. While the policy-maker's conversion to the Phillips curve concept is perhaps understandable (since he can be absolved from a failure to attain one of two desirable objectives), the role of the research economist in the specification of an appropriate inflation-unemployment policy has been remarkably passive. The simple theory of the Phillips curve, and its obvious policy implications, were quickly incorporated into the generally accepted corpus of macroeconomics. Clearly, if such a trade-off curve exists and is stable, one would expect economic research to be directed toward two issues: (1) determination of the precise economic costs of inflation and unemployment to form bases for an assessment of optimality amongst attainable combinations of inflation and unemployment, and (2) theoretical analyses of various policies designed to shift the trade-off curve to more favourable positions. Surprisingly, very little substantive analytical work has come forward on either count. This is particularly true with respect to the first issue,[1] and we are left with simplistic notions of debtor-creditor effects and lost output.[2] While recently there has been a lively empirical debate over the success of incomes policy, particularly in the United Kingdom,[3] theoretical analysis of policy options to shift the Phillips curve are infrequent.

On the other hand, analysts have been preoccupied with the empirical, and more recently the theoretical, foundations of the Phillips curve itself. During the 1960s a large outpouring of empirical work purported to demonstrate the unquestionable existence of the Phillips curve. Nonetheless, a strong theoretical attack, stressing the role of expectations, was launched by Friedman (1968) and Phelps (1967), and others, denying the theoretical existence of a stable trade-off curve. As Smith (1970) has illustrated, this theoretical dispute largely centers upon whether the coefficient for an expected price variable in the Phillips curve is equal to, or less than, unity.

Our purpose in writing this paper is to demonstate how sensitive the empirical estimates of the U.S. Phillips curve are to the underlying assumptions and estimation techniques currently being employed by wage analysts. It is our basic conclusion that the empirical evidence advanced to substantiate the claim that a stable Phillips curve exists for the U.S. is tenuous at best. This conclusion is not based on adjustments of time periods nor additions of new variables, although both of these strategies might substantiate our premise. Rather we accept both the economic theory and data employed by earlier contributions, and simply examine how sensitive these estimates are to underlying assumptions (implicitly) contained in their models.

Before turning to our own analysis, several comments on the general state of the arts in empirical Phillips-curve analysis are in order. As suggested above, there are a plethora of studies purporting to demonstrate the significance of unemployment and prices in determining money wage changes.[4] However, as one reviews this burgeoning literature, one cannot help but be astounded by the ingenious and pragmatic approaches of the Phillips-curve analysts. There appears to be an unending list of explanatory variables which have been found significant in various "augmented" Phillips curves. Perhaps even more disturbing is the intertemporal instability of the estimated coefficients, particularly the coefficient on the unemployment variable.[5] In short, studies which appear to have excellent statistical credentials (in inferential terms) have produced the following consensus on the Phillips curve:

(1) it clearly exists

(2) it is definitely not invariant to time nor author.

Paradoxically then, the greatest attack on the Phillips curve may implicitly rest in the diversity of the empirical evidence purported to prove its very existence!

Perhaps the only consistency in this perplexing maze of Phillips-curve estimates is the estimation of a particular functional form. With few exceptions, quarterly empirical wage studies have adopted the overlapping-annual-wage-change (OAWC) specification for the dependent variable, that is, either $(w_t - w_{t-4})/w_{t-4}$ or $(w_t - w_{t-4})$ where w_t is an aggregative index for wage levels and t is a temporal subscript, and fourth-order moving averages for the explanatory variables.[6] An incomplete list of studies which make use of these specifications would include the following works: Anderson (1969), Bodkin *et al.* (1966), de Menil (1969), Dicks-Mireaux and Dow (1959), Eckstein (1968), Evans and Klein (1968), Helliwell *et al* (1969), Kaliski (1972), Klein and Ball (1959), Kuh (1967), Levy (1967), Lipsey and Parkin (1970), Perry (1966, 1970), Phelps (1968), Pierson (1968), Reuber (1970), Schultze and Tryon (1965), Simler and Tella (1968), Vanderkamp (1972), and Vroman (1970).

Adequate reasons for the popularity of the OAWC model are somewhat difficult to ascertain. Usually it simply is adopted without any explanation. One might speculate whether this action is attributable to economists' general reluctance to justify the choice of functional form or to a belief that this particular model represents the existing paradigm and, therefore, requires no explanation. Infrequently, the model is briefly justified on the grounds that it will "avoid seasonality" and "reduce noise and measurement error" [Phelps (1968), p. 707]. However, more detailed accounts of the rationale and success of this unorthodox approach to these fundamental statistical problems are never given. Understanding of this problem of model selection and its inherent consequences can only be obtained by a review of certain institutional characteristics of the labour market.

As Tobin points out in his recent A.E.A. presidential address,

> "Keynes emphasized the institutional fact that wages are bargained and set in the monetary unit of account. Money wage rates are, to use an unKeynesian term, "administered prices". That is, they are not set and reset in daily auctions but posted and fixed for finite periods of time."
>
> [Tobin (1972), p. 3]

While Tobin discusses theoretical implications of such institutional arrangements, the implications for aggregation and estimation may be of equal importance. Recognition of this discontinuity of wage bargaining and reviews is quite explicit in the early studies of wage determination [e.g., Dicks-Mireaux and Dow (1959) and Perry (1966)]. In fact, the prevalence of such institutional features in the labour market prompted the adoption of the OAWC model with its fundamental distinction between discontinuous, unobservable, micro-wage relations for particular groups of the labour force and an aggregate relation formed from them. Given the proliferation of collective bargaining and the development of longer-term contractual agreements, the identification of such discontinuities and institutional features becomes even more important for the correct temporal specification of the explanatory variables[7] and for the implementation of an efficient estimation technique.

Since we have presented a complete analysis of the general model elsewhere,[8] only a brief statement of the aggregation-institutional assumptions sufficient for the specification of the OAWC are given. Most of these assumptions are either implicitly or explicitly stated in Dicks-Mireaux and Dow, and Perry.

Aggregation Assumptions for the Conventional Quarterly Wage Change Model

(A1) Wages are set annually for all workers, and, once established, remain fixed until the next annual negotiation and settlement.

(A2) The labour force is divided into four distinct groups on the basis of the quarter in which their annual wage negotiations and/or reviews take place.

(A3) The ratios of all seasonal groups in the labour force to the total labour force are constant. In other words, the percentage of workers who bargain in the j-th quarter of the year is constant over the entire sample period.

(A4) The percentage change in wages for each of the four seasonal groups is a function of the same set of explanatory variables with the same parameter values for each group. Explanatory variables (X) and error term (u) are dated in the quarter in which the wage negotiation, settlement, and/or review took place (j). That is,

$$\frac{w_j^h - w_{j-4}^h}{w_{j-4}^h} = aX_j + u_j$$

for h = 1, ... , 4 and where w_j^h is the wage-rate for the h-th group in the j-th quarter.

(A5) The relative change in the aggregate wage-rate is approximated by a moving average of the relative changes in the wage-rates for the four groups. The weights of this moving average are assumed to be equal (i.e. .25).[9] In essence this latter assumption equalizes the four seasonal bargaining groups.

$$\frac{w_t - w_{t-4}}{w_{t-4}} = \sum_{h=1}^{4} .25\left(\frac{w_t^h - w_{t-4}^h}{w_{t-4}^h}\right)$$

$$\frac{w_t - w_{t-4}}{w_{t-4}} = a\sum_{i=0}^{3} .25\, X_{t-i} + \sum_{i=0}^{3} .25\, u_{t-i} ,$$

where w_t is the aggregate wage-rate.

(A6) The micro error terms $\{u_j\}$ are assumed to be normally distributed with constant, equal variances.

The most important implication of these assumptions concerns the nature of the error term in the estimated equation. This error term is unequivocally autocorrelated since it is formed by a moving average of the underlying micro errors. The usual application of least squares to the OAWC model will yield biased estimates of standard errors and invalid t-statistics.[10] Appropriate estimation techniques are available; and, as we demonstrate below, their use reveals that the empirical and policy consequences of ignoring this inherent autocorrelation are substantial.

The empirical results presented in this paper are primarily based on the economic model advanced by Perry (1966). This is perhaps the best known of American studies. It has been employed by many other analysts, and we also examine an application by Pierson (1968) of the Perry model to assess the influence of union strength on the position of the Phillips curve. We explore the statistical effects of the inherent autocorrelation of the OAWC model, and the sensitivity of the estimates to the imposition of the set of institutional assumptions (given above), "convenient" assumptions which bear no resemblance to the real world of wage movements.

The Consequences of Inherent Autocorrelation

As we have reported elsewhere,[11] the OAWC model provides an ideal basis for the rare application of Aitken's technique of generalized least squares (GLS). Given that empiricists have had no difficulty in specifying the weights of the moving averages for the explanatory variables (i.e. 0.25 weights), the properties of the dispersion matrix for the error terms are easily obtainable since the same (known) set of weights is employed in both sets of moving averages. Table 1, reproduced from an earlier paper, [12] displays OLS and GLS estimates for Perry's basic model.[13] "While all variables are pseudo-significant under OLS, there is no consistent pattern for variables under GLS. No variable's coefficient retains significance in each of the three sectors. The basic Phillips-curve variable is significant in only one sector while price and profit variables are significant in two of three sectors. In sectoral terms, no one equation exhibits significant coefficients for all four explanatory variables."[14] In nearly all cases, the inherent autocorrelation of the OAWC model inflates the OLS t-statistics by more than 100% (usually 200-300%) over the appropriate GLS results.

In her study, Pierson extends Perry's model through 1966 Q2.[15] In Table 2, we present GLS estimates for all of the equations presented in her article.[16] As shown in the first row, the basic conclusions of Table 1 are reinforced. The model, as applied to total manufacturing, has only one significant variable, unemployment. Consumer price changes, profitability and the quidepost dummy are all insignificantly different from zero. To examine the effect of union strength, Pierson disaggregates total manufacturing in terms of 1958 unionization rates at the two-digit level. Groups I (strong) and II (weak) split the entire manufacturing sector into two exhaustive categories, while Groups A and B provide a greater divergence by selecting sub-groups at either end of the unionization rate spectrum.[17] In general, we note that:

- (i) consumer price changes are insignificant for either weak or strong bargaining groups,
- (ii) unemployment is insignificant except for Group I (strong union strength),
- (iii) lagged profits are significant for both weak union strength groups, but not for the strong union groups,
- (iv) change in profits is significant only for the most strongly unionized group, and
- (v) the guidepost dummy is only significant for the least strongly unionized group.

In short, nearly all of Pierson's arguments (e.g., "greater union strength is significantly associated with greater adaptation of wage changes to cost-of-living changes", p. 461) are either reversed or unsubstantiated when the inherent autocorrelation in the OAWC model is removed.[18]

As a further check on the differential effects of union strength on the Phillips curve, we have utilized a Chow test modified to correct for moving-average error terms.[19] To test the null hypothesis that Groups I and II, as well as A and B, have insignificantly different parameter estimates, calculated values for F-statistics are 3.94 and 1.26 respectively (in contrast to a critical .05 value of 2.19).[20] Thus, we conclude that two groups drawn from the extremes of the unionization rate spectrum (A and B) have insignificant differences in parameter estimates, but a simple exhaustive strong-weak unionization rate split reveals sig-

Table 1. PERRY Wage Equations for U.S. Manufacturing Industry (1948 II - 1960 III)

	Ordinary Least Squares			Generalized Least Squares		
	(1)	(2)	(3)	(4)	(5)	(6)
$\dot{C}_{t-1}$	0.385	0.327	0.518	0.183	0.267	0.313
	0.055	0.060	0.059	0.113	0.106	0.145
	6.946	5.487	8.717	1.614	2.553	2.151
$1/U_t$	14.611	10.344	14.601	12.795	6.506	9.352
	2.176	2.220	2.542	5.694	5.149	7.900
	6.716	4.659	5.744	2.247	1.263	1.183
R_{t-1}	0.434	0.524	0.223	0.544	0.521	0.283
	0.069	0.074	0.064	0.184	0.154	0.197
	6.270	7.058	3.511	2.954	3.365	1.437
ΔR_t	0.832	0.714	0.527	0.856	0.559	0.664
	0.176	0.202	0.156	0.279	0.311	0.277
	4.727	3.529	3.371	3.065	1.797	2.395
Const.	-4.421	-4.659	-2.171	-4.949	-3.657	-1.393
D.W.	1.188	0.811	1.358	2.19	2.32	2.06
F(5,44)	71.706	62.825	59.967	8.713	11.727	4.130

TABLE 2

GLS ESTIMATES FOR PIERSON STUDY

	Dep. Var.	Constant	$\dot{C}_{t-1}$	U_t^{-1}	ΔRi_t	Ri_{t-1}	G	$\dot{W}j_t$	D.W.	$\bar{R}^2$
(1)	$\dot{W}$	.4251	.04535 (.29)	14.5872 (2.68)	.4924 (1.31)		-.1765 (.52)		2.48	.16
(2)	$\dot{W}_I$	.4181	-.03544 (.25)	12.7279 (1.96)	.1512 (.45)	.06726 (.27)	-.2631 (.82)		2.09	.18
(3)	$\dot{W}_{II}$	-1.3120	.1209 (1.04)	6.9681 (1.67)	.1294 (.50)	.2637 (2.44)	-.2647 (.99)		2.19	.12
(2a)	$\dot{W}_I$	.9164	-.04468 (.33)	13.8897 (2.83)			-.2522 (.82)		2.14	.22
(3a)	$\dot{W}_{II}$	-1.0238		7.1287 (1.73)		.2520 (2.37)	-.2878 (1.11)		2.28	.14
(3b)	$\dot{W}_{II}$	-1.0374				.2929 (2.84)		.2677 (2.52)	2.29	.17
(4)	$\dot{W}_A$	1.7835	-.1061 (.47)	10.3122 (1.02)	.7964 (2.07)	.0117 (.04)	-.2073 (.41)		2.31	.10
(5)	$\dot{W}_B$	-0.7653	.07674 (.62)	3.3385 (.71)	.2929 (1.29)	.3804 (3.06)	-.6187 (2.02)		2.29	.18
(4a)	$\dot{W}_A$	1.8700	-.1066 (.48)	10.5484 (1.32)	.7906 (2.25)		-.2052 (.41)		2.31	.12
(5a)	$\dot{W}_B$	-0.3461				.3247 (2.90)		.1648 (2.19)	2.30	.17

where i denotes relevant disaggregation for profits
j denotes appropriate spillover variable.

nificant differences! Given such anomalous results we are hard pressed to conclude that "union strength does make a difference; it significantly worsens the trade-off between unemployment and inflation" (Pierson, p. 465). Rather it appears that a "little" union strength matters (78% contrasted to 53% unionization rates) but not a large dose of union strength (87% versus 45%). However, such tentative conclusions are clearly overshadowed by the dismal performance of the basic economic explanatory variables, as revealed by GLS t-statistics. A trade-off relationship only exists for one group (I), and not for the other three at all!

Variable Definition for the Perry Model

Dependent variable: annual percentage change in straight-time hourly earnings of production workers for total, durable and non-durable manufacturing,

$$\left[(W_t - W_{t-4}) / W_{t-4}\right].$$

$\dot{C}_{t-1}$: four quarter moving average of one quarter percentage change in the consumer price index

$$\left(\sum_{i=0}^{3} \frac{C_{t-i} - C_{t-1-i}}{C_{t-1-i}}\right),$$

lagged one quarter.

$1/U_t$: reciprocal of the four quarter moving average of the unemployment rate.

R_{t-1} : four quarter moving average of the annual profit rate (ratio of corporate earnings after taxes to stockholders equity), lagged one quarter, for total, durable and non-durable manufacturing.

ΔR_t : first difference of the profit rate series.

{(1), (4)} Total Manufacturing

{(2), (5)} Non-Durables Manufacturing

{(3), (6)} Durables Manufacturing

Autocorrelation at the Micro Level

The previous GLS estimates are based on the assumption that the underlying micro errors are normally distributed with constant, equal variances. To the extent that this assumption is violated, the GLS estimates will also be inefficient, and the concomitant statistical inferences invalid. In this section of the paper we examine the empirical sensitivity of the GLS estimates to various alternative distributions of the micro error terms.

In an earlier theoretical paper[21] we examine the role of autocorrelated micro errors in the context of an implicit wage spillover between various micro labour groups. The strategy adopted here is somewhat less ambitious and more empirically oriented. As a test of the empirical sensitivity of the OAWC model to various micro error distributions, we postulate several alternative "general" structures for these underlying error terms. Next, we conduct a number of experimental GLS regressions with various possible values for the parameters of the general structure. To the extent that statistical inferences and coefficient parametric estimates vary over these experimental regression sets, we draw conclusions concerning the sensitivity of the OAWC model to the specification of a particular distribution for the micro errors.

The two general structures for the micro error distributions which appear most plausible are the following:

$$\text{[A6a]} \quad u_t = \lambda u_{t-4} + v_t, \quad |\lambda| < 1,$$

and

$$\text{[A6b]} \quad u_t = \rho u_{t-1} + v_t, \quad |\rho| < 1,$$

where in both cases v_t is a stationary white noise sequence. The first structure is based on annual increments with an autoregressive structure internal to one particular micro labour group. The second structure provides an autoregressive structure between two successive micro labour groups. As mentioned above, this can be construed as an implicit wage spillover between bargaining groups and established the interdependence of all micro groups. In Tables 3 and 4, we present GLS estimates for the Perry economic model as applied to total manufacturing. In each case eight alternative autoregressive parameters can be compared to $\lambda = \rho = 0$, assumption [A6] the basis of Table 1. The empirical consequences of internal, fourth-order autoregressive autocorrelation are relatively minor. Statistical inferences are invariant with respect to the choice of the λ parameter and parametric estimates are reasonably stable (particularly for moderate λ values).

However, the GLS estimates incorporating external micro autocorrelation (i.e. implicit spillovers between groups), as shown in Table 4, necessitate a major re-assessment of the Perry model. Any positive autocorrelation between groups produces insignificant price and unemployment effects. These results were also obtained for both sub-sectors of manufacturing. Clearly the proposition that a significant trade-off curve exists in the Perry model depends critically on the assertion that the error term for workers receiving

TABLE 3

INTERNAL MICRO AUTOCORRELATION

TOTAL MANUFACTURING

	Prices		Unemployment		Profits		∆ Profits				
λ	Coef.	t-value	Coef.	t-value	Coef.	t-value	Coef.	t-value	Const.	D.W.	$\bar{R}^2$
+.8	.178	1.777	19.830	2.784	.707	3.428	.864	3.567	-7.959	2.37	.56
.6	.161	1.546	17.945	2.568	.724	3.480	.842	3.407	-8.111	2.35	.56
.4	.166	1.549	16.305	2.454	.681	3.382	.836	3.275	-7.281	2.33	.55
.2	.173	1.573	14.548	2.364	.619	3.219	.832	3.130	-6.182	2.29	.52
.0	.179	1.597	13.006	2.312	.546	3.001	.831	3.002	-5.001	2.23	.47
-.2	.181	1.599	11.889	2.307	.477	2.773	.834	2.908	-3.951	2.17	.40
-.4	.179	1.576	11.230	2.347	.422	2.578	.837	2.884	-3.164	2.11	.33
-.6	.174	1.532	10.945	2.415	.384	2.431	.835	2.797	-2.652	2.05	.27
-.8	.055	0.468	12.100	2.544	.330	1.943	.990	3.058	-2.226	1.62	.08

TABLE 4

EXTERNAL MICRO AUTOCORRELATION

TOTAL MANUFACTURING

	Prices		Unemployment		Profits		Δ Profits				
ρ	Coef.	t-value	Coef.	t-value	Coef.	t-value	Coef.	t-value	Const.	D.W.	$\bar{R}^2$
.8	-.077	.489	6.235	.563	.570	1.206	1.049	2.902	-3.230	2.67	.10
.6	-.032	.210	8.638	.876	.612	1.642	1.041	3.129	-4.325	2.61	.15
.4	.040	.283	11.082	1.342	.606	2.125	.987	3.186	-4.966	2.52	.24
.2	.116	.923	12.472	1.841	.576	2.578	.908	3.108	-5.096	2.40	.36
0	.179	1.597	13.006	2.312	.546	3.001	.831	3.002	-5.001	2.23	.47
-.2	.222	2.200	13.142	2.716	.524	3.379	.770	2.927	-4.874	2.04	.56
-.4	.249	2.665	13.146	3.024	.511	3.677	.730	2.891	-4.778	1.84	.62
-.6	.264	2.973	13.120	3.221	.504	3.872	.707	2.882	-4.718	1.70	.65
-.8	.273	3.144	13.095	3.322	.501	3.972	.697	2.885	-4.685	1.62	.66

(bargaining) wage increments in a particular quarter of the year bears *no* positive relationship to the error term of the previous group of wage bargains or reviews. Given the non-existence of spillover variables of any variety in the Perry model, such an assertion is somewhat dubious.[22]

Sensitivity to Institutional Labour Market Assumptions

Consider now the empirical validity and implications of the aggregation assumptions imposed. Unfortunately, data to examine these assumptions are primarily of a fragmentary nature. The evidence which does exist[23] strongly suggests that a large portion of the labour force bears no relationship to the standard set of assumptions. The presence of multi-year contracts, seasonal-bunching of wage settlements and reviews, and front-end loading, clearly contradicts the simple symmetric assumptions of the OAWC model.

Without venturing into a methodological debate about the realism of assumptions, we would point out two conclusions if such aggregation assumptions are invalid:

(i) Given the mis-specification of the moving-averages for the explanatory variables, the parametric estimates will be biased to the extent that these variables are serially correlated.

(ii) Since such mis-specification also directly affects the (moving-average) properties of the error term, classical statistical inferences will be inappropriate.

While bias in parametric estimates is a serious consequence of a mis-specification of the institutional-aggregation assumptions, the latter conclusion is perhaps more important. Explanatory variables may appear to be "significant" (and conversely variables may be incorrectly diagnosed as "insignificant") simply because the estimates of the variances are based on incorrect formulae. Furthermore, the aggregation-institutional assumptions cannot be verified through the successively passing of conventional statistical tests since only the "correct" aggregation assumptions will have unbiased estimates of variances. In short, one must assert that the aggregation assumptions are empirically valid and proceed with statistical inference to verify the theoretical relations postulated at the micro level.

Since the resulting Phillips-curve estimates have been used as an important basis in policy decisions and some of the underlying aggregation assumptions appear arbitrary, it is worthwhile to assess the sensitivity of estimated coefficients and their associated t-statistics to changes in the set of assumptions. Our preliminary analysis focuses on the conventional aggregation model presented above. All assumptions are retained with only one exception: the four micro labour groups are not constrained to be equal (A5). Thus, "seasonal bunching" is permitted in the context of the usual OAWC model with fourth order moving averages for all explanatory variables. Since there is no reason to believe that any one particular seasonal distribution of workers dominates another, we have simply drawn eight sets of random numbers for these distributions. In addition, four different permutations of each set are employed by varying the initial element in the sequence. An "equal" weights regression is calculated for comparative purposes.

GLS results for these thirty-three seasonal distributions are presented in Table 5.[24] As pointed out above, one cannot test the "significance" of the aggregation assumptions (in this case different seasonal bunching patterns). One must simply postulate an empirical set of aggregation assumptions as the basis for testing one's wage theory. Table 5 therefore provides 32 additional sets of estimates to compare with any one equation selected (by the reader?) to represent the closest approximation to the institutional features of the labour market. In other words, if one had stipulated a different seasonal pattern, would one have drawn different inferences concerning the significance and magnitude of various theoretical variables.

Clearly Table 5 reveals a high degree of sensitivity concerning parametric estimates for different seasonal bargaining patterns. In terms of significance levels, there is a dramatic difference between seasonal patterns 2 and 6 (four of five variables apparently significant at the .05 level) and patterns 8, 16, 21, 25, 26, 27 and 32 (none apparently significant even at the .10 level). In contrast to the usual equal-weight assumption, the postulation of almost any other seasonal bargaining assumption improves one's chances of detecting a significant coefficient for the major explanatory variables (CPI, R, and U). In short, the specification of the particular seasonal pattern has a pronounced effect on the statistical inferences drawn with respect to each of the explanatory variables. It would have been much more reassuring if variables were either consistently significant or insignificant across all seasonal patterns.

As pointed out above, variation in the estimated parametric values is expected since an inappropriate "seasonal bunching" assumption will introduce specification error and bias into the estimates. Estimated coefficients for the consumer price variable range from −.52 to +2.89 while estimated coefficients for the profit rate variable range from .45 to 5.12. Given the widespread interest in the position and shape of the Phillips curve, we have presented the thirty-three estimated Phillips curves in Sections A to H of Chart I.

TABLE 5

Sensitivity of Conventional OAWC Model to Specification of Seasonal Bargaining Distribution in U.S. Manufacturing Sector, 1953-68

	Seasonal Weight Set				Constant	$\dot{C}_{t-1}$	R_{t-1}	U_t^{-1}	G_t	F-test
(1)	.250	.250	.250	.250	-.0248	.2960	.4538	.0593	-.0172*	40.32
(2)	.106	.226	.243	.425	-.0592*	.6256	.6024*	.1184*	-.0218*	25.04
(3)	.425	.106	.226	.243	-.0379	-.1782	.6426*	.0444	-.0190*	20.75
(4)	.243	.425	.106	.226	-.0482	.1058	.6473	.0698	-.0227†	12.34
(5)	.226	.243	.425	.106	-.0216	.9552	.4902	.0300	-.0231*	15.41
(6)	.074	.229	.238	.459	-.0842*	.6582	.7428*	.1544*	-.0271*	16.36
(7)	.459	.074	.229	.238	-.0530	-.2117	.8204*	.0337	-.0217†	13.53
(8)	.238	.459	.074	.229	-.0693	.1163	.8235	.0769	-.0268	6.76
(9)	.229	.238	.459	.074	-.0204	1.0356†	.5293	.0143	-.0278†	8.68
(10)	.028	.458	.403	.111	-.2096	1.2088	1.5383	.3041*	-.0652*	5.32
(11)	.111	.028	.458	.403	-.1719*	-.4119	1.8310*	.0831	-.0408	4.06
(12)	.403	.111	.028	.458	-.1687	.0732	.6766	.4977†	-.0311	2.24
(13)	.458	.403	.111	.028	-.0521	2.8987†	1.3029	-.1657	-.0640	2.00
(14)	.008	.347	.395	.250	-.7347*	1.4985†	4.5834	1.0505*	-.1769†	4.28
(15)	.250	.008	.347	.395	-.4782*	-.5225	5.1204†	-.0497	-.0897	1.40
(16)	.395	.250	.008	.347	-.5619	.2200	3.1724	.9734	-.1044	.94
(17)	.347	.395	.250	.008	-.0155	2.4634*	2.0636	-.5040	-.1550	1.87
(18)	.415	.157	.388	.040	-.0219	1.2795*	.7838	-.0613	-.0431	3.53
(19)	.040	.415	.157	.388	-.1562*	.7488	1.0821	.2810*	-.0445	8.25
(20)	.388	.040	.415	.157	-.0940†	-.2558	1.3115	.0073	-.0319†	5.58
(21)	.157	.388	.040	.415	-.1327	.1097	1.0074	.2299	-.0332	3.09
(22)	.350	.340	.247	.063	-.0193	1.7857†	.6046	-.0269	-.0310†	6.72
(23)	.063	.350	.340	.247	-.0890*	.9761	.7890†	.1514*	-.0316*	13.25
(24)	.247	.063	.350	.340	-.0650†	-.3019	.8842*	.0513	.0222†	13.76
(25)	.340	.247	.063	.350	-.0719	.1142	.6462	.1569	.0227	5.80

TABLE 5 (Continued)

	Seasonal Weight Set				Constant	$\dot{C}_{t-1}$	R_{t-1}	U_t^{-1}	G_t	F-test
(26)	.456	.184	.100	.260	-.0466	-.0740	.5994	.0894	-.0202	10.66
(27)	.260	.456	.184	.100	-.0307	.7989	.5936	.0236	-.0264	11.00
(28)	.100	.260	.456	.184	-.0531†	.7349	.6393*	.0863†	-.0252	19.02
(29)	.184	.100	.260	.456	-.0541*	.1211	.6568*	.0877†	-.0192*	23.37
(30)	.097	.366	.170	.367	-.0660*	.4866	.6279†	.1339*	-.0243*	18.88
(31)	.367	.097	.366	.170	-.0373	-.0382	.6765*	.0321	-.0216	17.79
(32)	.170	.367	.097	.366	-.0607	.1264	.6197	.1223	-.0221	12.09
(33)	.366	.170	.367	.097	-.0217	.7850	.5482	.0158	-.0246*	12.32

* significant at .05 level

† significant at .10 level

CHART 1A

Phillips Curves for Different Seasonal Bargaining Patterns

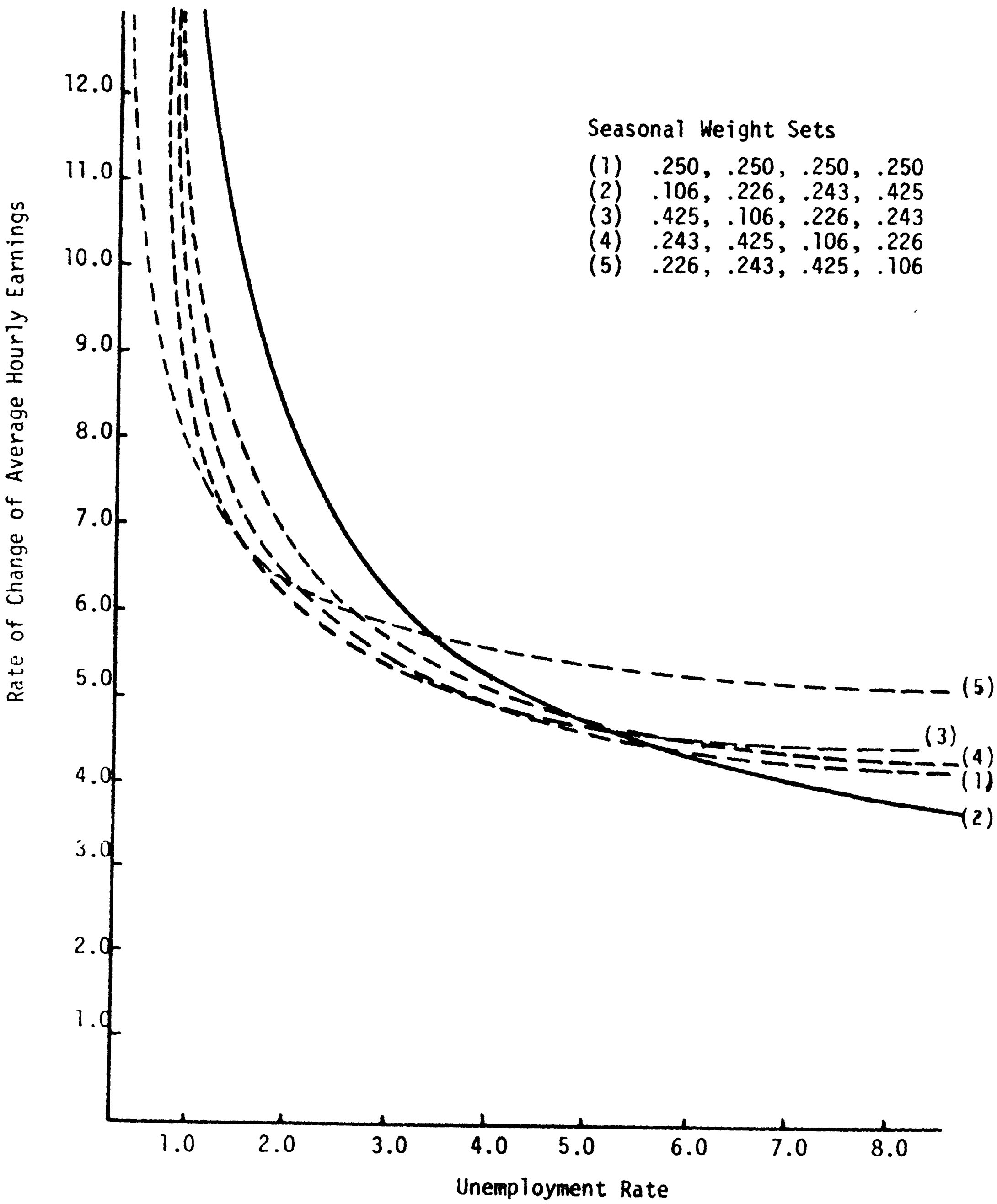


CHART 1B

Phillips Curves for Different Seasonal Bargaining Patterns

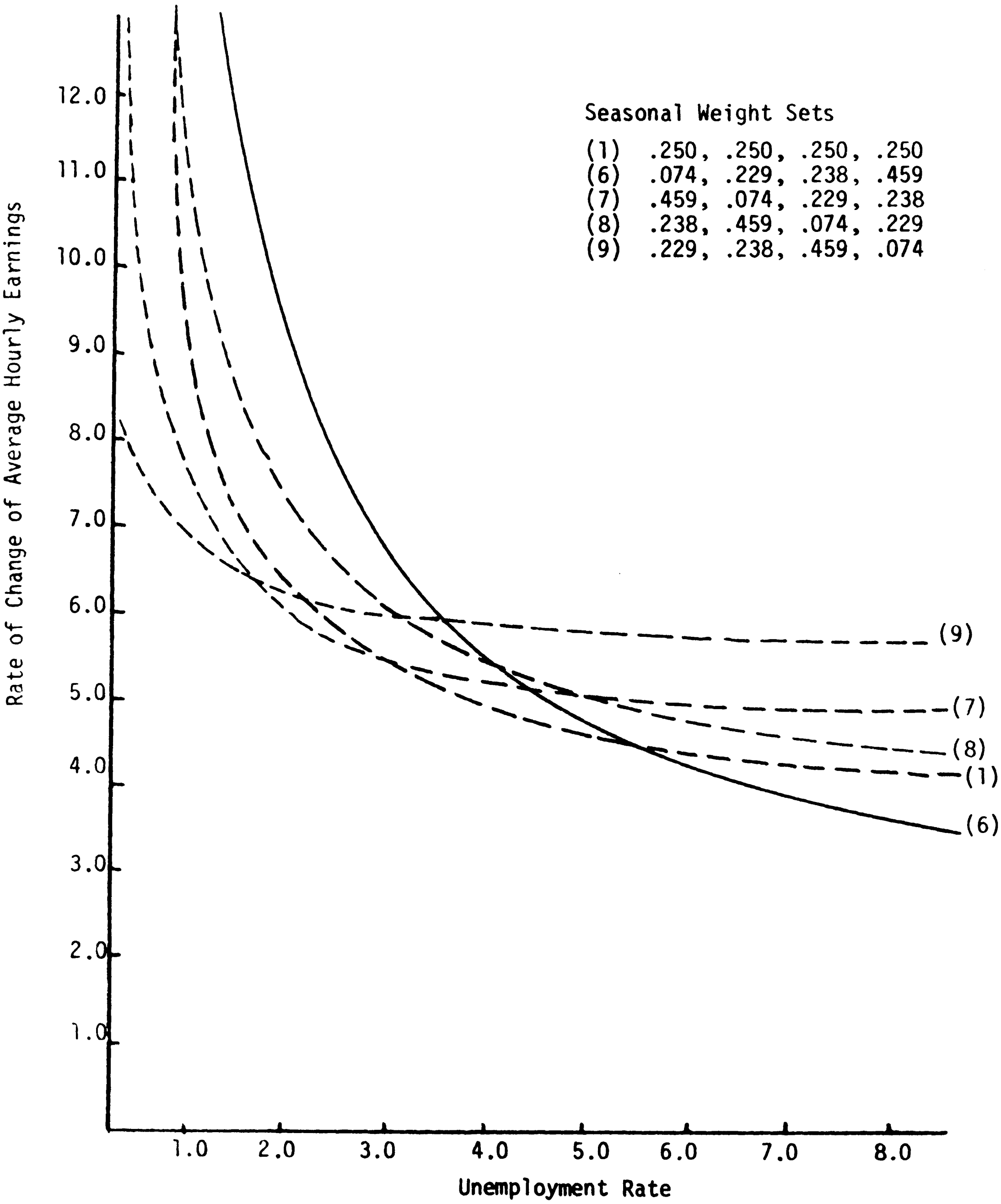


CHART 1C

Phillips Curves for Different Seasonal Bargaining Patterns

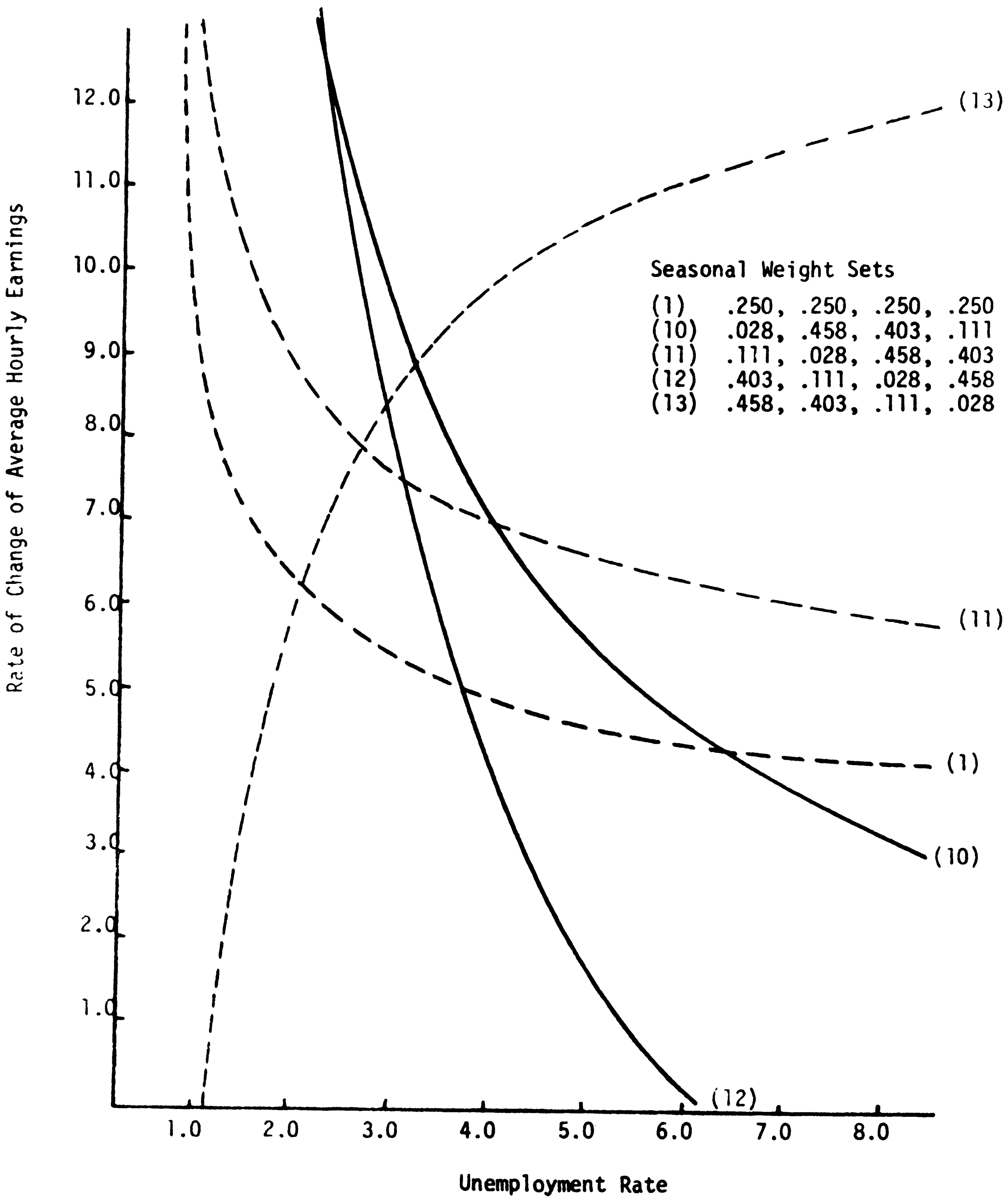


CHART 1D

Phillips Curves for Different Seasonal

Bargaining Patterns

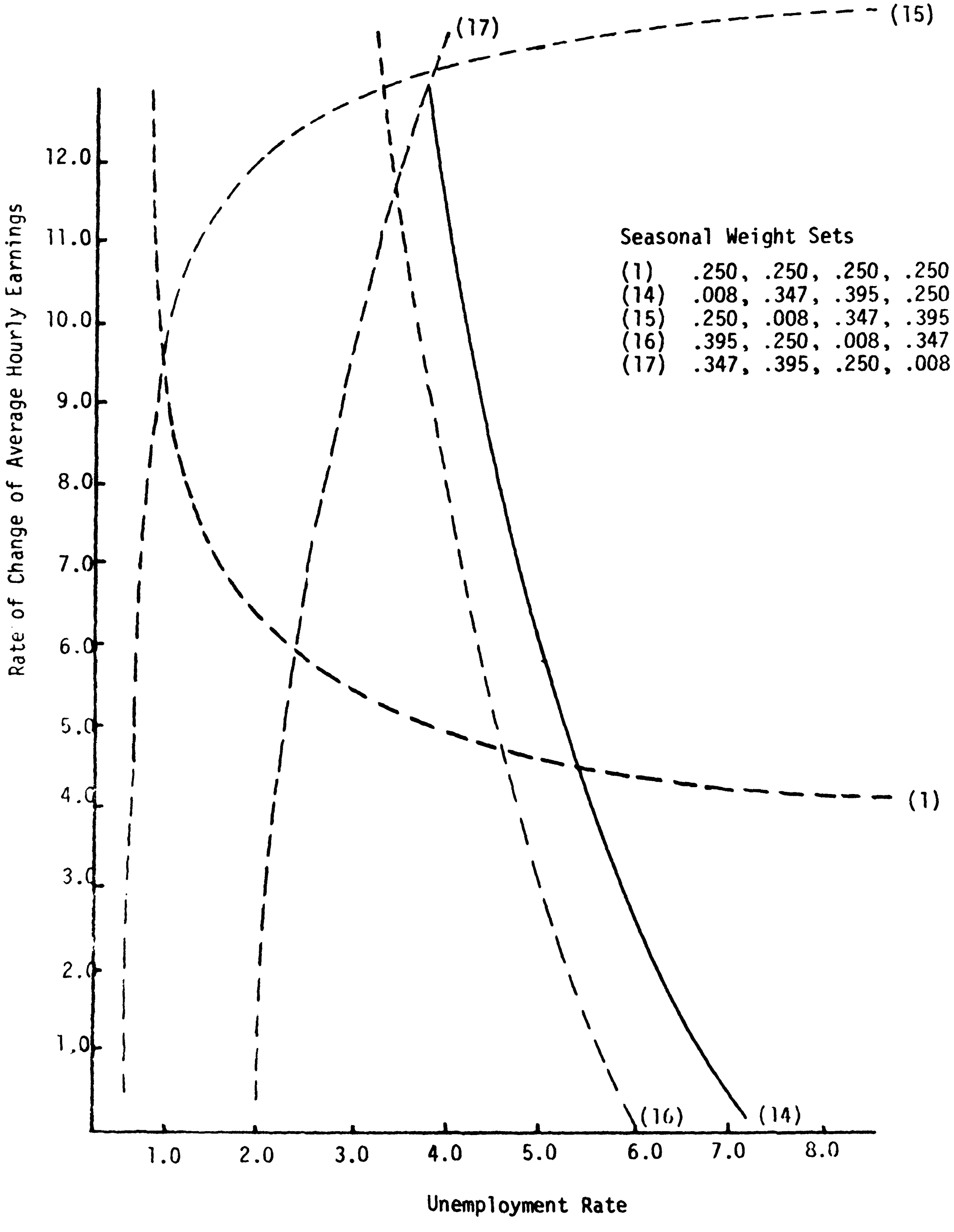


CHART 1E

Phillips Curves for Different Seasonal Bargaining Patterns

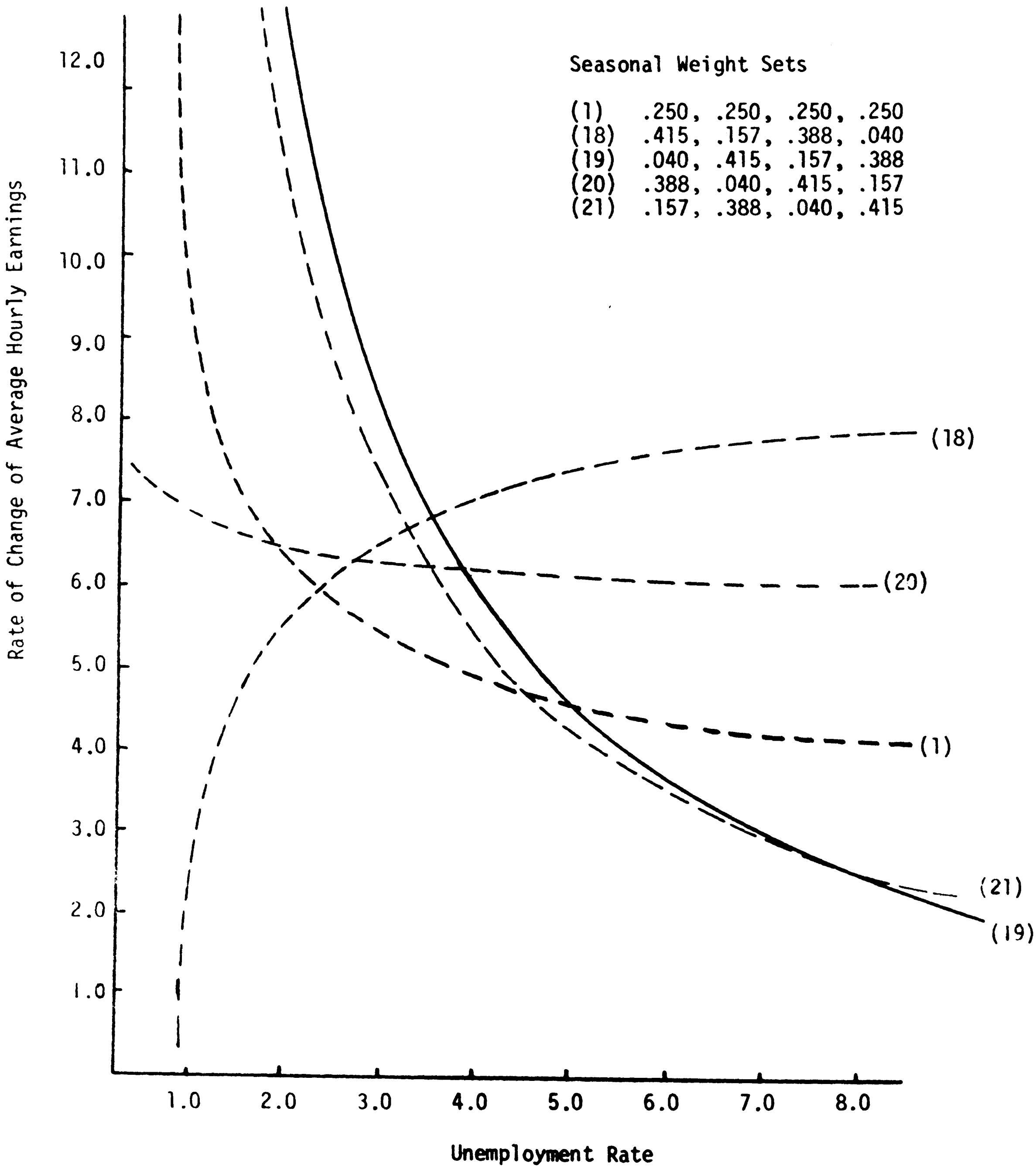


CHART 1F

Phillips Curves for Different Seasonal Bargaining Patterns

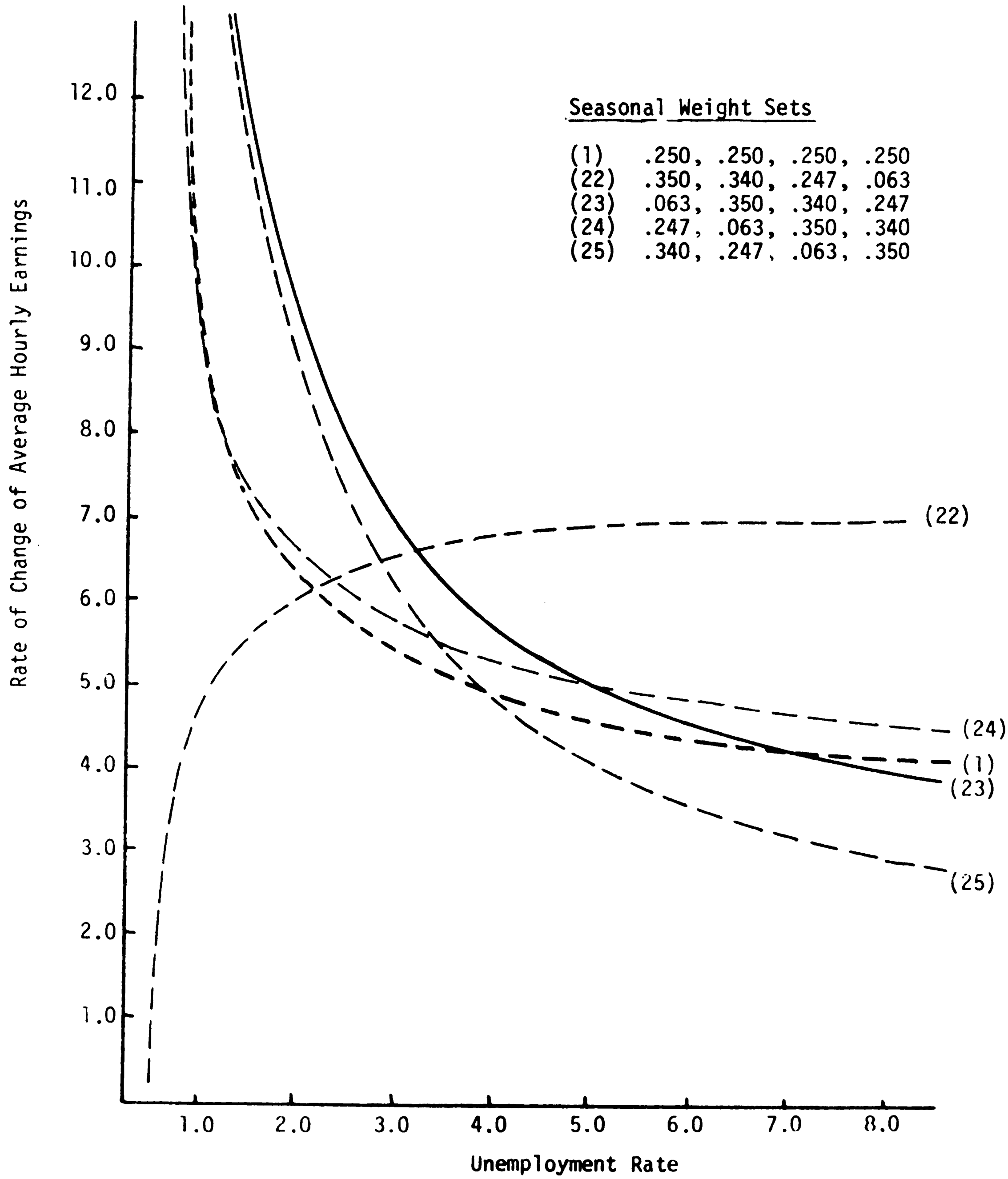


CHART 1G

Phillips Curves for Different Seasonal Bargaining Patterns

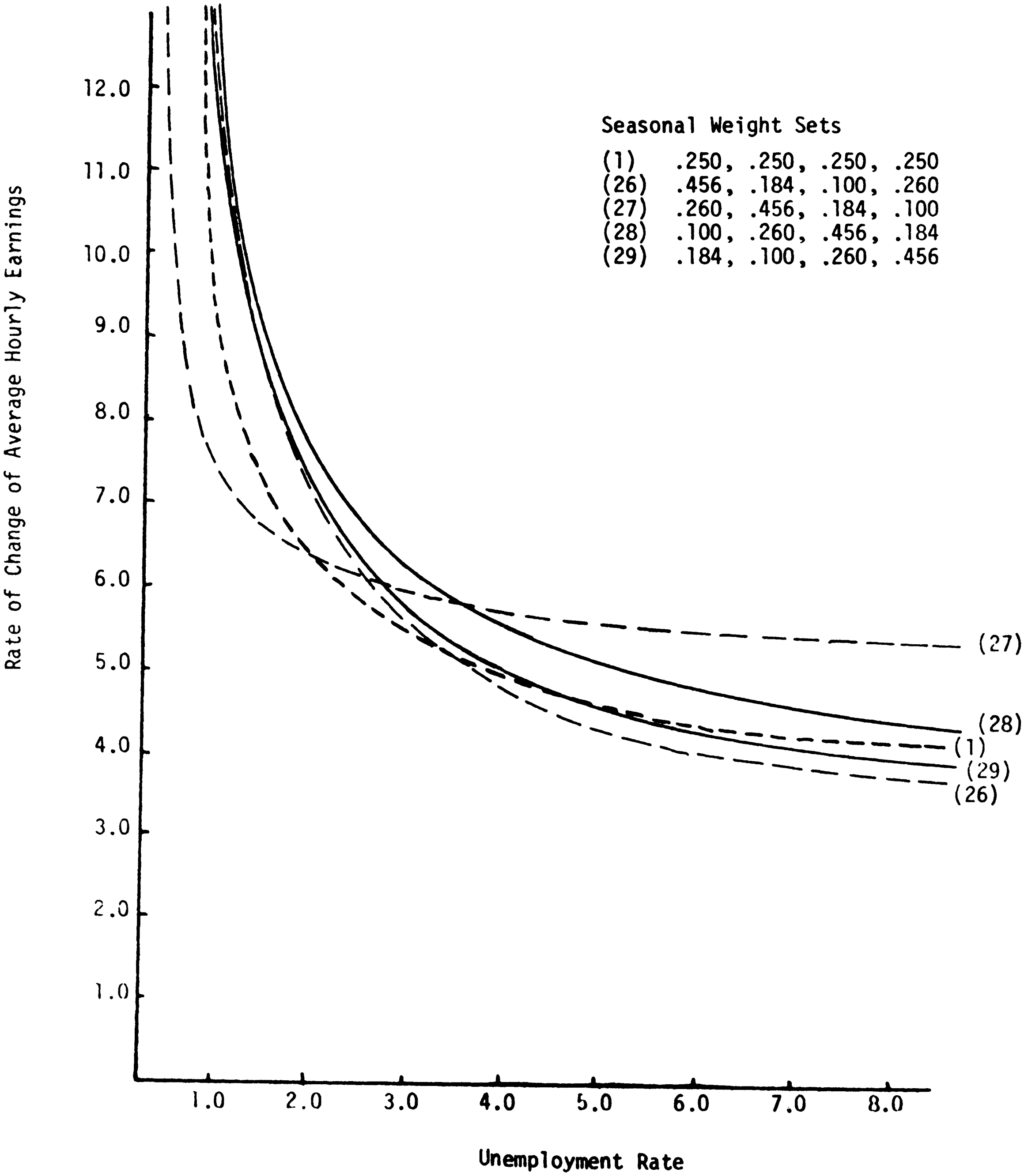
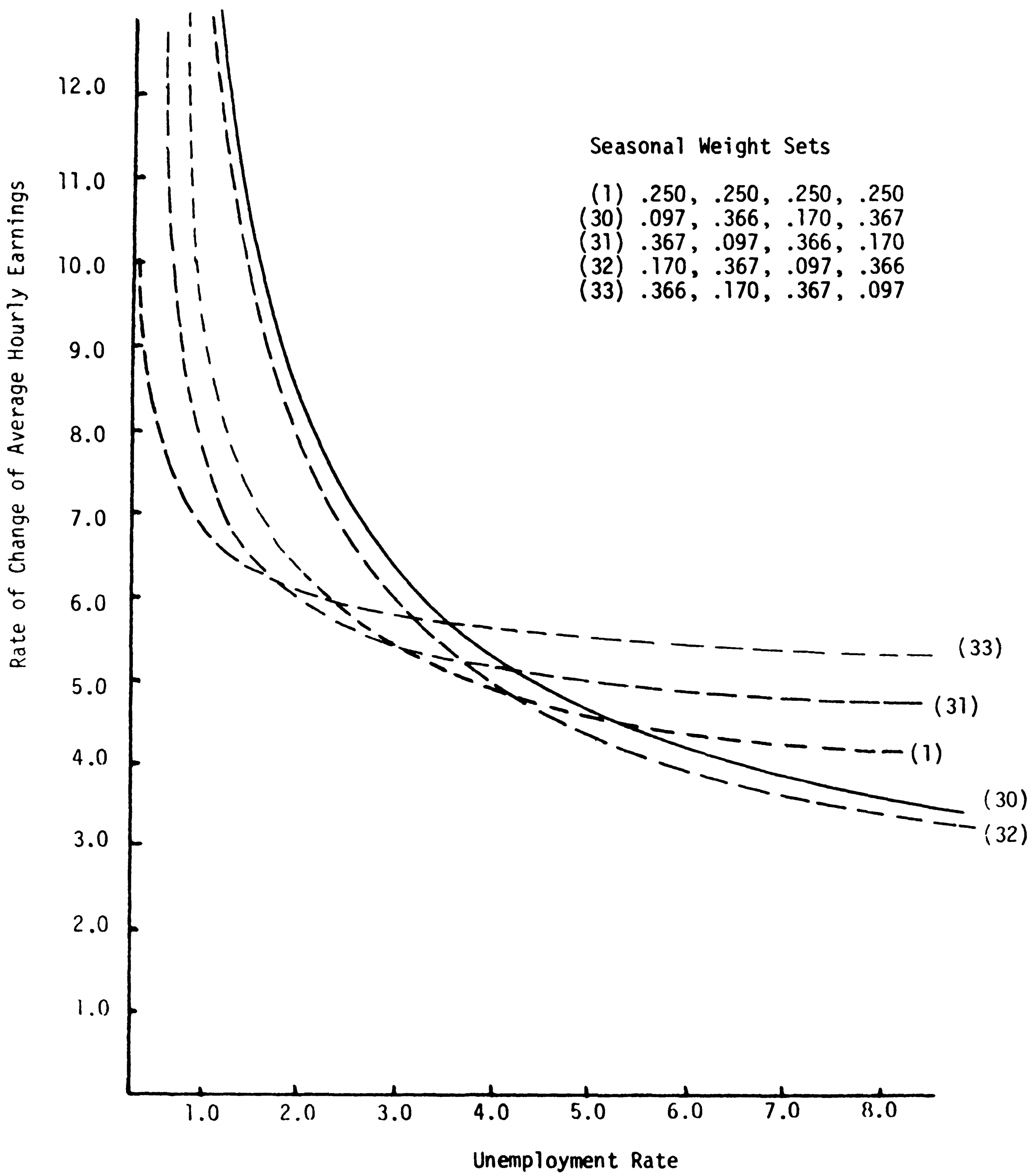


CHART 1H

Phillips Curves for Different Seasonal

Bargaining Patterns



(All other variables are specified in terms of their 1968 levels). Each Section of the Chart depicts the four permutations of a particular weight set together with the "equal weight" Phillips curves (the conventional estimates). A solid line, as opposed to a broken line, signifies (apparent) significance at the .10 level. Again, there is considerable variation depending upon which weight set one selects to represent the seasonal pattern of wage bargains and/or reviews. In particular, one notes the dramatic differences in the Phillips curves presented in Charts 1C, 1D and 1E. Even if one restricts oneself to apparently significant estimates, the Phillips curves generated with weight sets (10), (12), (14) and (19) are substantially more inelastic than those generated with weight sets (2), (6), (23), (28), (29) and (30).

In a previous paper[25] we have generalized the OAWC model to incorporate multi-year contracts, front-end loading and seasonal bunching. A large number (143) of regressions were calculated making various assumptions concerning these three institutional features of the labour market. As a summary device, histograms are presented for these 720 estimated coefficients in Chart II. The results conform to those obtained in the previous experiment. Besides the diversity in parametric estimates, statistical inferences (the shaded areas represent significance at the .10 level) drawn from different sets of aggregation-institutional assumptions are remarkably unstable.

Conclusion

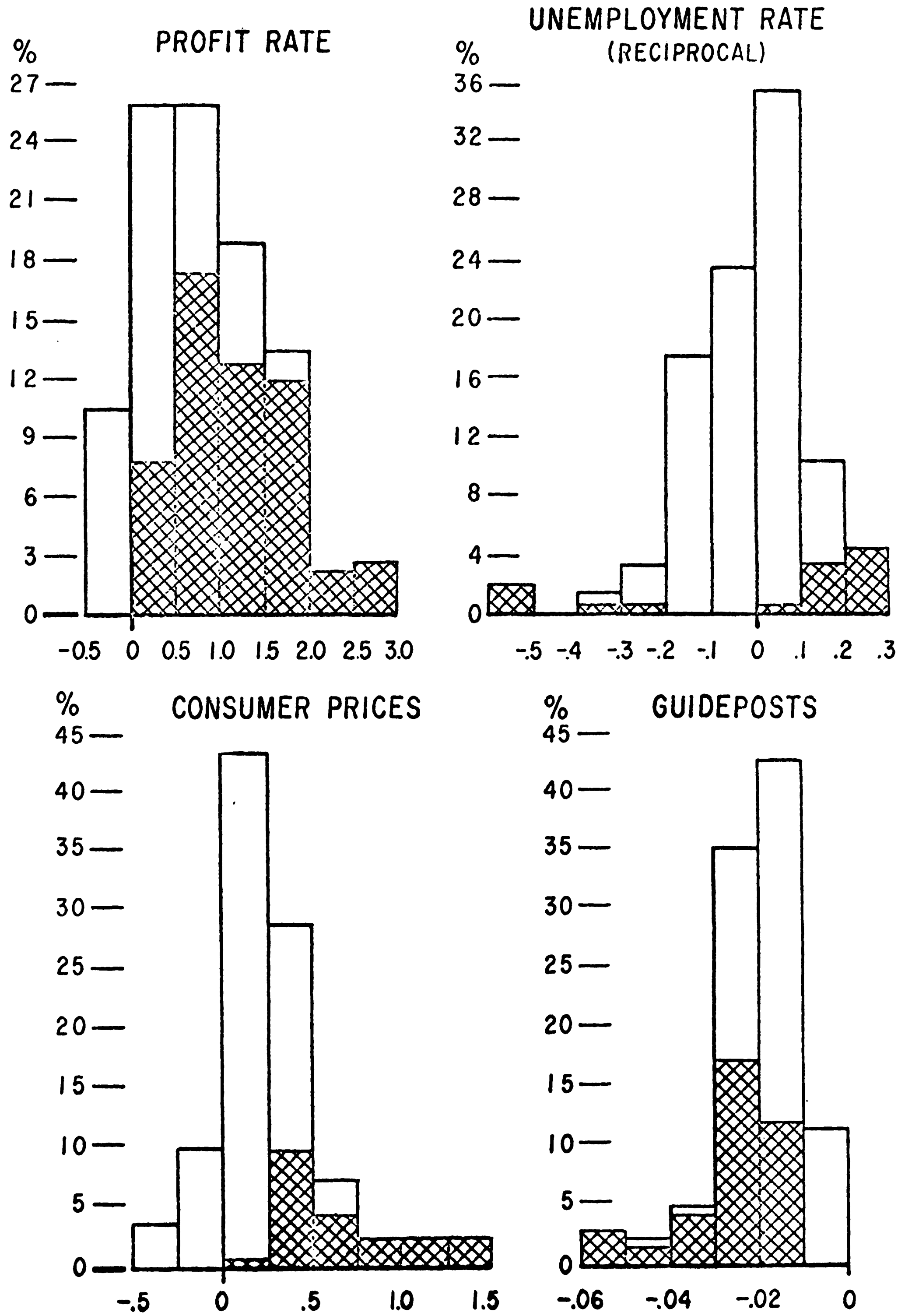
The adoption of the OAWC model is based on a particular set of "institutional," labour market assumptions. Autocorrelation in the error term is a direct consequence of the aggregative procedure indicated by the assumptions. In particular, if researchers specify the same moving average properties for all explanatory variables, then the aggregate error must exhibit the same moving-average properties. Criticism of previous empirical work which fails to correct for this problem is not a minor cavil. Efficient estimates of wage equations presented by Perry and Pierson indicate that many of the coefficients for hypothetical explanatory variables are insignificantly different from zero.

Relaxing the assumption of independently distributed micro errors appears to further aggravate the problem. While GLS estimates are rather insensitive to the existence of fourth-order, within-group, micro-autocorrelation; the presence of positive first-order micro autocorrelation between successive micro groups substantially lowers GLS calculated t-statistics. In short, inherent (uncorrected) moving-average autocorrelation in the OAWC model coupled with even a small amount of between group micro autocorrelation sufficiently inflates OLS t-statistics to produce erroneous inferences concerning the significance of the critical economic variables in quarterly wage equations.

Finally, such estimates are highly sensitive to the imposition of institutional labour market assumptions with respect to multi-year contracts, front-end loading, and seasonal bunching. The diversity of the parametric estimates and the apparent contradictory statistical inferences contained in both experiments strongly indicate that the institutional details of the labour market must be given far more attention than they have received in recent quarterly wage research. Further, the empirical patterns suggested by the various sets of institutional-aggregation assumptions are sufficiently perverse for us to assert that no simple correctional device will necessarily reduce the effects of our ignorance of these critical institutional features. Statistical analysis of the OAWC cannot reveal the nature of the wage determination process without a substantial effort by both users and collectors of economic data to clarify the structural form of the institutional background in the labour market. Until such an effort takes place, the general consensus that a stable Phillips curve exists, and has a particular shape, is based on questionable empirical foundations.

CHART II

HISTOGRAMS OF ESTIMATED COEFFICIENTS



FOOTNOTES

1. The Hollister and Palmer (1969) study is an exception. Also, an excellent survey is provided by Foster (1972).
2. Even this issue appears to be more of a distributional problem than generally is recognized since a disproportionate portion of additional output is received by those who find employment.
3. See, for example, the conference report edited by Johnson and Nobay (1972) which provides many criticisms of the empirical results of Lipsey and Parkin (1970). In the mid- to late sixties, a somewhat milder debate emerged in the U. S. concerning the Kennedy-Johnson guideposts [see, for example, Anderson (1969)]. One anticipates another "rash" of dummy variables for the recent Nixon policies to emerge soon in U. S. empirical studies.
4. Such studies are also quite unanimous in rejecting the Friedman-Phelps effect.
5. For example, Levy (1967) reports the following unemployment coefficients for four consecutive subperiods in the 1948-67 time period: 34, 23, 60, and 8.
6. Recent studies by Eckstein and Wyss (1971) and Gordon (1972) employ centered changes of various dimensions. These pragmatic specifications are subject to the same criticism as presented in the text of this paper.
7. Laidler (1972) indicates the crucial role of the institutional features in the generation of price expectations.
8. Rowley and Wilton (1971).
9. If this assumption is only approximately correct, other methods of estimation might be considered. Some results embodying stochastic parameters are contained in Rowley, Smith and Wilton (1972). If this assumption is exact, then movements of the wage index are excessively restricted. See Rowley and Wilton (1973a).
10. See Rowley (1972a).
11. Rowley and Wilton (1973b).
12. *Ibid.*
13. Data for these regressions was taken from the appendices of Perry (1966).
14. Rowley and Wilton (1973b). For two Canadian studies, GLS results revealed *no* significant explanatory variables.
15. It should be noted that Pierson has dropped the first four and one-half years of the Perry data span and added a guidepost dummy variable.
16. We wish to thank Gail Pierson for readily making her data available to us. Since our OLS results are almost identical to hers, we have suppressed them from this table.
17. Average unionization rates for the four groups are 78 percent, 53 percent, 87 percent, and 45 percent. For further details, see Pierson (1968), p. 457-8.
18. Unlike most OAWC model analysts, Pierson is aware of the problems of this positive autocorrelation (see first paragraph, p. 461), but clearly unaware of the severity of the problem.
19. For further details on this modification, see Rowley and Wilton (1972a).
20. OLS Chow tests produce calculated F-values of 4.58 and 3.44 respectively.
21. Rowley and Wilton (1972d).
22. In a study in progress, spillovers from other recent wage settlements appear to be as important as the role of more traditional variables in the determination of money wage rate changes.
23. See for example, L. A. Dicks-Mireaux and J. C. R. Dow, *op. cit.*, U. S. Department of Labour, B.L.S. Reports 102 and 282 and Bulletin 1353. Also, see recent evidence collected by Rowley and Wilton (1972e) for the Canadian economy.
24. One notes the insignificance of all explanatory variables except the Guidepost dummy in this data set based on Perry (1970).
25. Rowley and Wilton (1972c).

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