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I. Introduction

The old controversy about whether wage and price controls are instrumental or not in bringing down the rate of inflation was never quite buried in the U.S., although after the apparent failure of Nixon's Economic Stabilization Program of 1971-1974 (ESP) to achieve its desired goals, even the enthusiasm of Incomes Policy advocates appeared for a while to have been mitigated. Defeat however, was never acknowledged by its defenders. All type of ex-post rationalizations were offered as to, why, in the case of controls having been properly implemented and coordinated with more responsible monetary and fiscal policy, they would have eventually succeeded and not been thwarted in holding the reins of an inflation which from a brisk trot before the imposition of the ESP had set forth in an unbridled gallop at the end of it.

Vanquishing inflation has certainly never been a pure academic problem since the costs of either living with it or putting an end to it have always been too real for everybody. But the Reagan administration's belief in a supply-side free lunch as an effective weapon against inflation has rekindled the debate. There is an increased willingness to have yet another go at controls of those who always seeing wage inflation as the big culprit do not think that moderate fiscal and monetary restraint together with the measures to reduce costs and boost productivity (as Reagan's administration has set out to do) can alone curb price inflation without the parallel complementary effect of some direct mandatory wage restraint.

The main opponents of policed wage-setting have always been the hard-core monetarists who very rightly pointed out the inability of "cost-pushers" to distinguish between absolute and relative price increases when they use a concept such as "monopoly-power" to explain how generalized inflation can be set off. If inflation is a monetary phenomenon, everywhere, so their argument goes, controls are ineffective since they attack the symptoms without doing anything about the cause. The only point on which both main rival groups agreed was that affecting expectations of inflation in a downward direction would do the trick. However the way to go about it signals yet another point of departure beween them since one can always argue that if controls do not seem to affect inflation, and were unsuccessful during the past in fighting it, why people would continue to believe they will now succeed in doing so.

On the other had the recent experience of Mrs. Thatcher's government in the U.K. is showing that just applying steady monetary and fiscal squeeze on demand to curb inflation (even with parallel preamouncement of monetary policy to influence inflationary expectations) may be a much slower and painful remedy than what originally was made out to be (necessary creation of the so called "reserve army of unemployed").

It will not be attempted here to give a detailed description of all the arguments for and against controls as a remedy against inflation, since it would extend the length of the study far too much and it has already been thoroughly studied and documented in the literature, (for a good overview see for instance: Andersen and Turner (1), Jack Carr (5). The purpose here is limited - to take again the skeleton out of the cupboard, namely the U.S. Economic Stabilization Program of 1971-1974 and try to ask some questions similar to the previous quantitative studies about its effects, as well as some other questions which were perhaps not so much looked into before.

II. OVERVIEW OF THE E.S.P.

For a variety of reasons not so important here, the Federal Reserve System's monetary policy from the early sixties to the imposition of controls by President Nixon on August 15, 1971 was basically to inject money into the economy at an accelerating pace; it comes then as no surprise that the trend rate of inflation also accelerated. Those responsible for the printing presses though, did not exactly assumed responsibility for the discomforting outcome of it all and the well-loved myths of monopoly power either on the side of business or labour, to explain undefeatable inflation were nurtured by businessmen, unionists and politicians alike in an atmosphere of self-delusion. Retrospectively it comes then as no surprise that when the controls program was announced the only mandatory restrictive ceiling was that on increases in nominal wages in unionized firms - in plain words, an attempt at meddling with the monopoly power of unions, with the sins of wages.

Following the main lines of a much more detailed description of the four Phases of the E.S.P. program by Darby (6.) I shall here give a summary of their main characteristics.

Phase I: 1971,9 - 1971,11

It was argued that the E.S.P. could speed the adjustment to a lower rate of inflation by reducing inflationary expectations and revising labour contracts since the wage demands of nonunionized workers and the agreed wages in union contracts included an adjustment for expected inflation. The temporary decrease in employment linked to a reduced growth rate in the money supply could thus be reduced or obliterated. This initial freeze on nominal wages was conceived as a surprise attack in order to avoid anticipatory strategic wage and price increases while a more sophisticated controls program was being prepared. The price freeze restricted most prices just trivially if at all.

Phase II: 1971,12 - 1972,12

Prices were allowed to increase proportionately to increases in costs, but this profit margin rule was a binding ceiling only where there were major shifts in relative demand or supply. The rate of inflation could then only be affected by the program if the rate at which costs were growing was checked by wage controls which is what characterized Phase II.

Nearly all unionized firms could still hire all the labour they wanted under the controlled wage rate and an incentive to evade the wage controls due to fear of losing employees did not actually existed. The main effect of controls on union wages was to increase the number of workers which unionized firms were ready to employ.

Prior approval to raise prices was required, but firms which in normal circumstances, would have increased their prices by a certain amount, were able to meet the smaller legal ceiling by making up the difference by reducing the quality of the products so as to reduce their unit costs.

Therefore the reported price indices was an underestimate of the real ones since they did not allow for the drop in product quality.

Phase III: 1973,1 - 1973,6

Prior approval of price increases except in a few industries was removed for large firms although the profit margin of the previous phases remained binding as did the previous 5.5% annualized standard for nominal wage increases. As a result, quoting Darby (6), these firms were free to increase their prices not only in proportion to the increase in costs required to produce a unit of given quality but also in proportion to the increased costs necessary to restore the previous degraded quality. Since this restoration in quality, was not taken into account, the reported rate of inflation was an overestimate, although the difference between the reported price level and the "true" one was not completely reduced. As a result the annualized rate of inflation in these first six months of 1973 looked six percent points higher than in the last six months of 1972 and Phase III was stopped to give way to the next and last period of the E.S.P.

Phase IV: 1973,7 - 1974,4

A freeze on prices was initially attempted at the beginning of this period with nominal wages, however, allowed to rise as before. Shortages and illegal evasion of the controls became then the rule as firms were confronted with frozen prices lower than their market level. After only two month the previous notification and approval of price increases which had characterized Phase II was reinstated. The unit profit limit was again largely ignored but apparently binding enough so as to cause in the end the E.S.P. supporters on the business side to join the unions in calling for an end of the controls.

III. QUANTIFYING THE EFFECTS OF WAGE AND PRICE CONTROLS

It is evident that the effect of controls cannot simply be gauged by a direct comparison of the rates of wage and price change before the imposition of controls and the rates of change during them. Other factors affecting the labour market, like for instance a change in fiscal or monetary policy may have influenced the path of wage - price inflation. To properly assess the effect of controls one needs an economic model explaining the normal process of wage and price determination. Standard practice has always been to generate forecasts with this estimated model over the period of controls and compare the actual observed values of the relevant variables with the predicted ones to give an estimate of the "shock" on the normal development of variables such as wages and prices resulting from the enforcement of controls.

A common procedure has been to use structural models of wage-price determination, normally some variant of the expectations augmented Phillips curve relationships, based on a trade-off between inflation and unemployment. For instance

$$\hat{W} = f_1(\hat{P}^e, \bar{U}, \hat{\Pi})$$
$$\hat{P} = f_2(\hat{W}, \hat{\Pi}, \hat{PI})$$

where

$\hat{W}$ = rate of growth of money wages
$\hat{P}$ = rate of change in prices
$\bar{U}$ = unemployment rate
$\hat{\Pi}$ = rate of growth of labour productivity
$\hat{PI}$ = rate of change of import prices
$\hat{P}^e$ = expected rate of inflation

In relation to the endogenous variables in a model such as the last one, the policy impact is normally captured by using the estimated reduced forms. This procedure of course introduces an element of arbitrariness since the values of the "exogenous" variables which are necessary to generate counter factual predictions of the endogenous variables during the period of controls may have themselves been contaminated by the controls program. Recent research carried aout by Sims (25) rejects the hypothesis that, variables such as $\bar{U}$ and GNP can be considered exogenous im a "structural" model similar to the one above.

For instance one could argue that import prices $\hat{PI}$ in a system of flexible exchange rates, will be affected by controls in the case of inflationary expectation being actually revised down by controls.

Productivity growth also may be affected by increased friction in labour markets. Even the observed unemployment rate during the controls may be partly affected by them if the initial apparent success in bringing down the rates of wage and price inflation gives ground to the government to believe that they have effectively pushed down the Phillips curve southwesterly and tempts policy makers into increasing the rate of money expansion by more than they should. This last type of perverse effect seems to have been the case in the U.S. after the initial apparent success of Phase I.

A further problem of this traditional approach even assuming that the structural equations are properly specified and suitable estimated (which very often has not been the case, see Oi (17)), is the assumption of structural stability which may be a strong assumption to make if anticipation of and reaction to wage-price controls have altered the structure of the wage-price equations in the periods preceding and following the controls period, since the observations corresponding to these periods are normally used to fit the "normal" structure imposed on "policy-off" observations.

I cannot extend myself here to give an exhausting review of the main econometric criticisms of the numerous empirical studies using these type of simulation techniques and/or intercept dummy variables to model the impact of incomes policy (the interested reader should refer to Oi (17), although some of these difficulties referred to, and quite some more will always be present when attempting a quantification of the "shock" of controls by the above mentioned or alternative techniques like the ones used in this paper.

The approach used here is the identification, fitting and further use for generating counterfactual predictions of several variables,of time series models of the ARIMA type developed by Box-Jenkins. The rationale for doing so is explained in more detail in section IV.

An identical methodological approach has been used by Feige and Pearce (8) to also investigate the impact of the E.S.P. on several variables, albeit, only during Phase I and Phase II of the E.S.P. However their study concentrated on the effect upon the Consumer Price Index, Wholesale Price Index and Average Hourly Earnings for the whole private nonfarm economy.

On the other hand in this paper only the manufacturing sector is considered since after all, the manufacturing sector in the U.S.A. is heavily unionized and as we have already seen in Section II, controls were aimed first and foremost at the unionized firms. Furthermore in Feige and Pearce study, no considerations about the effect of controls on real wages were taken into account, since they did not estimate directly a real wage generating process like it is done here and the only conclusions one could draw from their study about real wages were of an indirect nature by comparing relative movements in wages and prices. A procedure which in their study must be looked at with suspicion, since they used seasonally adjusted data to fit the wage equation doing away with one of the main advantages of ARIMA models which is their ability to describe "seasonality", "trend" and "residual" series under the umbrella of one overall model.

A reason to emphasize directly real wage effects should be clear from the description of the several Phases in section II. Not only the impact on wages and prices may have been asymmetrical but also the underreported rate of inflation as a result of a drop in product quality is an interesting point on its own. The neoclassical theory of employment determination, postulates a negative relationship between real wages and employment with changes in real wages preceding changes in employment as we know.

It would be impossible for any government to monitor any policy of direct meddling with the real wage through controls as a means to increase employment without their statistical agencies allowing for changes in product quality to construct a price index which would really be the one upon which firms take their decisions to either hire or fire workers.

A conscious attempt at increasing employment through decreased real wages does not seem to have been the stated aim of the economists who tailored the E.S.P., although if decreases in real wages occurred during the E.S.P. accompanied by increases in employment like some evidence presented here seems to point out to, it would also be a meaningful axercise to attempt some preliminary investigation of how the real wage-employment nexus, if any, was affected by controls.

IV. USE AND JUSTIFICATION OF ARIMA MODELS

The starting point of the Box and Jenkins approach is the assumption that the observations from a time series y_t are realizations from a stochastic process. As proved by Wold (21), for any stationary stochastic process there exists a linear decomposition in the form of a weighted sum of random shocks ε_t

$$y_t = \psi(B)\varepsilon_t \qquad (1)$$

where $\psi(B) = 1+\psi_1 B+\psi_2 B^2+\psi_3 B^3+\ldots..$

and B is the lag operator $B^i\varepsilon_t = B\varepsilon_{t-i}$ and the random shocks ε_t satisfy

$$E(\varepsilon_t) = 0, \quad E(\varepsilon_t\varepsilon_{t-s}) = \sigma_\varepsilon^2 I.$$

In order to transform equation (1) into a more amenable form for estimation, Box and Jenkins developed a parsimonious general class of time series models known as autoregressive integrated moving average models (ARIMA) and represented in their simplest form as

$$\emptyset(B)\Delta^d y_t = \Theta(B)\varepsilon_t, \qquad \Delta^d = (1 - B)^d$$

where $\emptyset(B) = 1 - \emptyset_1 B - \emptyset_2 B^2 - \ldots - \emptyset_p B^p$ is a p^{th} order polynomial with roots outside the unit circle, $\Delta^d y_t$ is the d^{th} difference of the series y_t necessary to achieve stationarity and finally $\Theta(B)$ is a q^{th} order polynomial with roots outside the unit circle.

One can then say that time series y_t has a representation as an ARIMA (p, d, q) process which not only is it able to yield a model like (1) but also, if certain invertibility conditions are met (see Box & Jenkins), an autoregressive process of infinite order of the form

$$\phi(B)\theta^{-1}(B)\Delta^{d}y_t = \varepsilon_t \qquad (2)$$

or $\Pi(B)\Delta y_t^{d} = \varepsilon_t$

An appealing feature of ARIMA models is that they can be interpreted as adaptive or error learning forecasting models which include as a subset the well known adaptive expectations models of the Koyck and Cagan sort. To bring this point out, one can represent any stationary series like

$$y_t = \Psi(B)\varepsilon_t = \varepsilon_t + \sum_{i=1}^{\infty}\Psi_i\varepsilon_{t-i}$$

and use as a forecasting criterion that the forecast of y_t for a lead time ℓ, $\hat{y}_{t+\ell}$, minimizes the expected mean square forecast error

$$E_t\{[y_{t+\ell}-\hat{y}_{t+\ell}]^2\} = E[e_t^2(\ell)] \qquad (3)$$

If we now use a set of weights Ψ_i^* to define the optimal forecast $\hat{y}_t(\ell)$ which minimizes $E[e_t^2(\ell)]$, we can write

$\hat{y}_t(\ell) = \sum_{j=0}^{\infty}\Psi_{\ell+j}^{*}\varepsilon_{t-j}$, weighted average of those error terms which we <u>can</u> estimate

Therefore the expected mean square forecast error at time t is

$$E[e_t^2(\ell)] = \sum_{i=0}^{\ell-1}\Psi_i^2\sigma_\varepsilon^2 + \sum_{j=0}^{\infty}(\Psi_{\ell+j}-\Psi_{\ell+j}^{*})^2\sigma_\varepsilon^2$$

which is minimized for $\Psi_{\ell+j}^{*} = \Psi_{\ell+j}$, $\forall j$

that is $E[e_t^2(\ell)] = \sum_{i=0}^{\ell-1}\Psi_i^2\sigma_\varepsilon^2 \qquad (4)$

Thus the minimum expected mean square forecast of $y_{t+\ell}$ is written as

$$\hat{y}_t(\ell) = \sum_{j=o}^{\infty} \Psi_{\ell+j} \varepsilon_{t-j} \qquad (5)$$

which can be seen to be the conditional expectation of $y_{t+\ell}$ at time t on present and past shocks

$$\hat{y}_t(\ell) = E \mid y_{t+\ell} \mid \varepsilon_t, \varepsilon_{t-1} \cdots\cdots \mid =$$

$$= E \mid \sum_{i=o}^{\ell-1} \Psi_i \varepsilon_{t+\ell-i} + \sum_{j=o}^{\infty} \Psi_{\ell+j} \varepsilon_{t-j} \mid \varepsilon_t, \varepsilon_{t-1} \cdots\cdots \mid =$$

$$= \sum_{j=o}^{\infty} \Psi_{\ell+j} \varepsilon_{t-j}$$

since $E[\varepsilon_{t+j}] = o$ for $j > o$

Similarly the ℓ-step ahead forecast made at time

t+1 is $\hat{y}_{t+1}(\ell) = \sum_{j=o}^{\infty} \Psi_{\ell+j} \varepsilon_{t-j+1}$

and since $\hat{y}_t(\ell+1) = \sum_{j=o}^{\infty} \Psi_{\ell+j+1} \varepsilon_{t-j}$

we obtain $\hat{y}_{t+1}(\ell) = \hat{y}_t(\ell+1) + \Psi_\ell \varepsilon_{t+1}$ (6)

which tells us, the expected value of the series y_t in the future, namely, the period $t+\ell+1$ is revised when new information is received.

Moreover the random shocks driving the stochastic process y_t can be interpreted as the one-step ahead forecast errors

$$e_t(1) = y_{t+1} - \hat{y}_t(1) = \Psi_o \varepsilon_{t+1} = \varepsilon_{t+1} \qquad (7)$$

since $\Psi_o = 1$ as can be seen by equating coefficients on different powers of B in equation

$$y_t = \frac{\Theta(B)}{\phi(B)} \varepsilon_t = \Psi(B) \varepsilon_t$$

So the economic unit at time t+1, revises all his previous expectations of future values of the series y_t by the amount $\Psi_\ell \varepsilon_{t+1}$ after he experiences an error, ε_{t+1}, in his previous forecast of y_t for t+1.

Now it can be easily seen that the well known adaptive expectations model of Koyck and Cagan (4.) is a very particular example of this class of models since it can be written as

$$\Delta y_t = (1-\theta_1 B)\varepsilon_t$$

or $$y_t = \Psi(B)\varepsilon_t$$

where $\Psi(B) = \frac{(1-\theta_1 B)}{(1-B)}$ giving $\Psi_i = 1-\theta$, $\forall i$, which means that as result of using the same weights Ψ's, the economic unit forecasts the same value of y_t for all future time periods.

Thus, refering back to the general class of ARIMA models, they can be said to use optimally all the predictive information contained in the past history of a series.

At a first superficial glance one might easily dismiss ARIMA models as rather "naive" forecasting tools when compared to bigger structural econometric models, since an ARIMA appears seemingly to explain the development of a series over time ignoring prevailing economic theory which would take into account all the possible influences and feedback of other relevant variables, which might be used to impose a particular structure on the data to generate a more powerful forecasting tool.

That this criticism is in principle valid is undeniable but fortunately can be qualified to a great extent since it turns out that the effect of other variables upon the series y_t we are attempting to forecast may already be contained in the past history of that series. As shown by Zellner and Palm (22) the ARIMA models, represent the "final form" for a variable implied by a very sophisticated model indeed. They show that a multivariate time series process will under very general assumptions always yield a set of processes generating individual variables, which have an autoregressive-moving average (ARMA) form.

To emphasize this extremely important point, let us use as an illustration a simplified version of the expectations augmented Phillips curve often used to justify incomes policies

$$\hat{W}_t = a_o + a_1 \hat{P}^e_t + a_2 f(U_t) + a_3 \hat{\Pi}_t + e_{1t} \qquad (8)$$

$$\hat{P}_t = b_o + b_1 \hat{W}_t - b_2 \hat{\Pi}_t + b_3 \hat{PI}_t + e_{2t}$$

$$\hat{P}^e_t = \hat{P}_{t-1}$$

Where

$\hat{W}_t$ = rate of growth of money wages

$\hat{\Pi}_t$ = rate of growth in labour productivity

$\hat{P}_t$ = rate of change of prices

$\hat{PI}_t$ = rate of change of import prices

$\hat{P}^e_t$ = expected rate of price inflation

e_{1t}, e_{2t} = error terms

If we now consider the vector

$Z'_t = (\hat{W}_t, \hat{P}_t, f(U_t), \hat{\Pi}_t, \hat{PI}_t)$ as being generated by a multiple time series process, it can be represented in matrix form as

$$H(B) Z_t = F(B) \varepsilon_t \qquad t=1, \ldots\ldots T \qquad (9)$$

where $\varepsilon'_t = (\varepsilon_{1t}, \varepsilon_{2t}, \ldots \varepsilon_{5t})$ is a vector of random errors such that $E(\varepsilon_t) = 0$, $E(\varepsilon_t \varepsilon'_{t+s}) = \delta_{t,t+s} I \quad \forall t,s$

And the matrices H(B), F(B) are both 5x5 matrices assumed nonsingular, whose elements are finite polynomials in the lag operator B and we denote as

$$H_{ij}(B) \quad \text{and} \quad F_{ij}(B) \quad \text{where } i,j=1, \ldots 5$$

Now what economic theory does is to impose "prior" information of exogeneity in the form

$$H_{ij}(B) = 0, \quad F_{ij}(B) = 0$$

for some i,j elements, to identify a structural model like (8) which can be used to generate causal equations known as reduced forms which relate functionally the current value of a variable such as $\hat{W}_t$ to the lagged endogenous and current and lagged exogenous variables and have traditionally been used in the econometric literature using simulation and dummy variables to assess empirically the effect of incomes control upon certain policy target variables such as wage or price inflation (Gordon (10), Eckstein and Brinner (7), Perry (19), Lipsey & Parkin (16)).

If we now go back to the representation (9) of the dynamic simultaneous equation model

$$H(B) Z_t = F(B) \varepsilon_t$$

we can write $Z_t = H^{-1}(B) F(B) \varepsilon_t$

or $$Z_t = \frac{H^{Adj}(B)}{|H(B)|} \cdot F(B) \varepsilon_t$$

where $H^{adj.}(B)$ = adjoint matrix of H(B)

$|H(B)|$ = determinant of H(B) which of course is a scalar, finite polynomial in B.

If the invertibility condition, $|H(B)| = 0$ has roots outside the unit circle, is satisfied we can express each variable in our vector

$$Z'_t = (\hat{W}_t, \hat{P}_t, f(U_t), \hat{\Pi}_t, \hat{PI}_t)$$

as a finite order ARMA equations, since

$$|H(B)|Z_t = H^{Adj.}(B)F(B)\varepsilon_t$$

or $|H(B)|Z_t = \Theta(B)\varepsilon_t$ (10)

where $\Theta(B) = H^{Adj.}(B)F(B)\varepsilon_t$

and Z_{it} for $i = 1,2,\ldots 5$ can be represented as $|H(B)|Z_{it} = \theta_i'\varepsilon_t$
$i = 1,\ldots 5$
where $\theta_i' = (\theta_{i1}, \theta_{i2}, \ldots \theta_{i5})$ is the i^{th} row of the Θ matrix.

Note now that $\theta_i'\varepsilon_t$ is a sum of moving average processes which can be represented as a moving average process in a single random variable.

Thus the ARMA processes for individual variables like the ones in vector Z_t' may be consistent with the "true" model dictating the behaviour of these variables,* and as a result they are suitable for forecasting purposes albeit not so for structural analysis. As shown by Nelson (14), expectations based on these univariate ARIMA models will in general have a larger mean squared error than expectations which are rational in the sense that they employ all the available information in the model, but as pointed out by Feige & Pearce (8) one must take into account the information costs incurred in the process of forming rational expectations. If we assume that the information cost for an economic unit to acquire information about the past history of a variable is relatively small, that is an added bonus to legitimate the use of ARIMA models as a clever forecasting tool.

To conclude, the identification estimation and further use for forecasting of univariate ARIMA models can be seen as an

*(which may not necessarily be one like (8)).

astute way of opening up carefully the Pandora's box that the real world is and let the data inside tell us an interesting and hopefully informative tale.

On the other hand, the imposition of sometimes too much arbitrary "a priori" theory and unsuitable econometric methodology to enable us to accept or reject the imposed structure (see Granger & Newbold (11) for study of nonsense regressions and spurious fits all-pervasive in the econometric literature), may be tantamount at times to a rash opening of the box's lid and remembering the good old myth, by so doing, Hope alone remained inside the box when all objects of desire were scattered to play havoc among mankind.

V. ESTIMATED ARIMA MODELS

The starting point is then to identify and estimate the ARIMA models most suitable to describe the underlying process generating the time series we are interested in. The identification procedure involves a detailed examination of the sample autocorrelation and partial autocorrelation functions which will point towards a suitable degree of differencing (parameter d) to achieve stationarity and the order p and q of the autoregressive and moving average terms respectively, (see Box and Jenkins (2)). Once the process has been initially identified, we choose that set of parameter values which will minimize the sum of squared residuals

$$S(\hat{\emptyset},\hat{\theta}) = \sum_t \varepsilon_t^2$$

where $\varepsilon_t = \theta^{-1}(B)\emptyset(B)\Delta^d y_t$

and $\hat{\emptyset}' = (\hat{\emptyset}_1,\ldots\hat{\emptyset}_p)$, $\hat{\theta}' = (\hat{\theta}_1,\ldots\hat{\theta}_p)$

since by assumption the error terms $\varepsilon_1 \ldots \varepsilon_t$ are all normally distributed and independent with mean 0 and variance σ_ε^2, their joint density is

$$f(\varepsilon_1,\varepsilon_2 \ldots \varepsilon_\tau) = \sigma_\varepsilon^{-n}(2\Pi)^{\frac{-n}{2}} \cdot e^{-\sum_{t=1}^{\tau}(\varepsilon_t^2/2\sigma_\varepsilon^2)}$$

So that the conditional log likelihood function associated with the parameter values $(\emptyset,\theta,\sigma_\varepsilon)$ is given by

$$L(\emptyset,\Theta,\sigma_\varepsilon) = -T \log \sigma_\varepsilon - \frac{\sum_t \varepsilon_t^2}{2\sigma_\varepsilon^2}$$

As we then can see, minimizing the sum of squared residuals $S(\emptyset,\Theta) = \sum_t \varepsilon_t^2$ with respect to the parameters $\emptyset,\Theta$ gives us their maximum likelihood estimates when the errors are assumed to be normally distributed. The minimization of the function $S(\emptyset,\Theta)$ which is nonlinear in the parameters, requires the application of iterative methods to achieve that set of parameters satisfying a preestablished convergence requirement.

One an ARIMA model has been fitted, the next step in assessing its adequacy is the checking of whether or not the residuals ε_t seem to be white noise as indeed they must be if the model is suitable. This is normally done by examining the estimated residuals correlogram

$$\hat{r}_k = \frac{\sum_t \hat{\varepsilon}_t \hat{\varepsilon}_{t-k}}{\sum_t \hat{\varepsilon}_t^2}$$

which under the hypothesis of serial independence should be themselves uncorrelated, normally distributed random variables with mean O on variance 1/T.

Complementary also, a "portmanteau"test devised by Box and Pierce follows immediately by summing up several individual squared residual sample autocorrelating $\hat{r}_k$ to obtain the approximate test statistic

$$Q = T\sum_1^k \hat{r}_k^2 \sim \chi^2(k-p-q)$$

under the hypothesis of the residuals being white noise.

These different steps of identification, estimation and diagnostic checking were applied using monthly data (since the timing and length of the different phases of the ESP makes it desirable) for the period 1953,7 - 1971,8, prior to the imposition of controls for the following variables, all expressed as annualized monthly rates of change (see Appendix for a description of the data).

CPI : consumer price index
WPI : Wholesale Price Index (Manufacturing)
WHE : Hourly earnings adjusted for overtime (Manufacturing)
EM : Manufacturing sector employment
RWC : Consumer Real Wage
RWW : Producer Real Wage

The results are reported in Table 1 where as can be seen some seasonal ARIMA models of the form $\phi(B)\Phi(B^s)y_t^d = \theta(B)\Theta(B^s)\varepsilon_t$ were found suitable.

The diagnostic checking of the residuals $\hat{\varepsilon}$ with the exception of those obtained from the CPI model (quite high Q statistic) was considered generally satisfactory and the models reported in Table I were used to generate simulated values over the ESP different periods to be compared with the reported actual values of the different series and thus obtain a quantitative estimate of the impact of controls.

TABLE 1 Estimated ARIMA models
maximum likelihood estimates and associated standard errors - Monthly data: 1953,7 - 1971,8

CPI :	$(1-B)y_t = \underset{(0.03)}{(1-0.84B)}\varepsilon_t$ $Q(41) = 49.4 \qquad \hat{\sigma}_\varepsilon = 2.37$
WPI :	$(1-B)y_t = \underset{(0.049)}{(1-0.69B)}\varepsilon_t$ $Q(41) = 30.42 \qquad \hat{\sigma}_\varepsilon = 3$
WHE :	$(1-B^{12})y_t = \underset{(0.06)}{(1-0.21B)}\underset{(0.05)}{(1-0.66B^{12})}\varepsilon_t$ $Q(40) = 24.98 \qquad \hat{\sigma}_\varepsilon = 5.22$
EM :	$(1-B)(1-B^{12})y_t = \underset{(0.05)}{(1-0.70B)}\underset{(0.04)}{(1-0.8B^{12})}\varepsilon_t$ $Q(40) = 24.60 \qquad \hat{\sigma}_\varepsilon = 12.36$
RWC :	$(1-B^{12})y_t = \underset{(0.07)}{(1-0.23B)}\underset{(0.04)}{(1-0.77B^{12})}\varepsilon_t$ $Q(39) = 40.2 \qquad \hat{\sigma}_\varepsilon = 5.58$
RWW :	$(1-B^{12})y_t = \underset{(0.07)}{(1-0.28B)}\underset{(0.04)}{(1-0.77B^{12})}\varepsilon_t$ $Q(40) = 32 \qquad \hat{\sigma}_\varepsilon = 5.47$

VI. MODEL SIMULATION DURING THE ESP

Some of the caveats related to using wage-price structural models to generate counterfactual predictions during the period of controls which have already been mentioned in section III, apply as well to the methods described here. If models such as those reported in Table I are supposed to characterise the processes generating those variables and if the imposition of controls shocks these normal processes, we can then generate optimal ARIMA forecasts from the models in Table I to be compared with the actual data during the controls. But would the ARIMA simulated values reproduce faithfully the path of those variables which would have taken place without controls?

In order to do so they would have ideally to take into account all the shocks affecting these variables in the absence of controls.

Therefore, generating updated, one-step ahead forecasts would be ideally suited since they are continously updated based on post sample observations. However this post-sample observations could be really misobservations, when the actual shocks do not match the hypothetical ones. Unfortunately, there is no available methodology, as mentioned in Feige and Pearce (9) to relate hypothetical with actual shocks and therefore it makes sense to use our estimated ARIMA models to generate also deterministic, ℓ-step ahead forecasts during the period of controls. Since these ℓ-step ahead forecasts assume that all future random shocks take on their expected value of zero and are based only on sample period data, they are not affected by events in the period of controls, 1971,9 - 1974,4.

Both deterministic and updated simulated values were calculated using the estimated models of Table I and compared with actual post-sample realizations. The results for the four phases of the ESP are reported in Tables 2, 3, 4, 5, 6, 7 and plotted in Figures 1 to 6.

The reported standard deviations of forecast error reported in the Tables make use of the fact that

$$\text{var}\left[\sum_{i=1}^{\ell} a_i e_t(i)\right] = \sum_{i=1}^{\ell} a_i^2 \, \text{var}\left[e_t(i)\right] +$$

$$+ \sum_{i \neq j} \sum a_i a_j \text{cov}\left[e_t(i) e_t(j)\right]$$

where a is a constant (in our case $1/\ell$). We have already seen in section IV that

$$E\left[e_t^2(\ell)\right] = \sum_{i=o}^{\ell-1} \Psi_i^2 \sigma_\varepsilon^2$$

And one can also derive immediately the following expression for the covariances

$$E\left[e_t(\ell) e_t(\ell+j)\right] = \sigma_\varepsilon^2 \sum_{i=o}^{\ell-1} \Psi_i \Psi_{j+i}$$

Since by definition the updated forecasts are uncorrelated, we take into account the covariances only for the computation of the standard deviation of the deterministic ℓ-step ahead forecast errors.

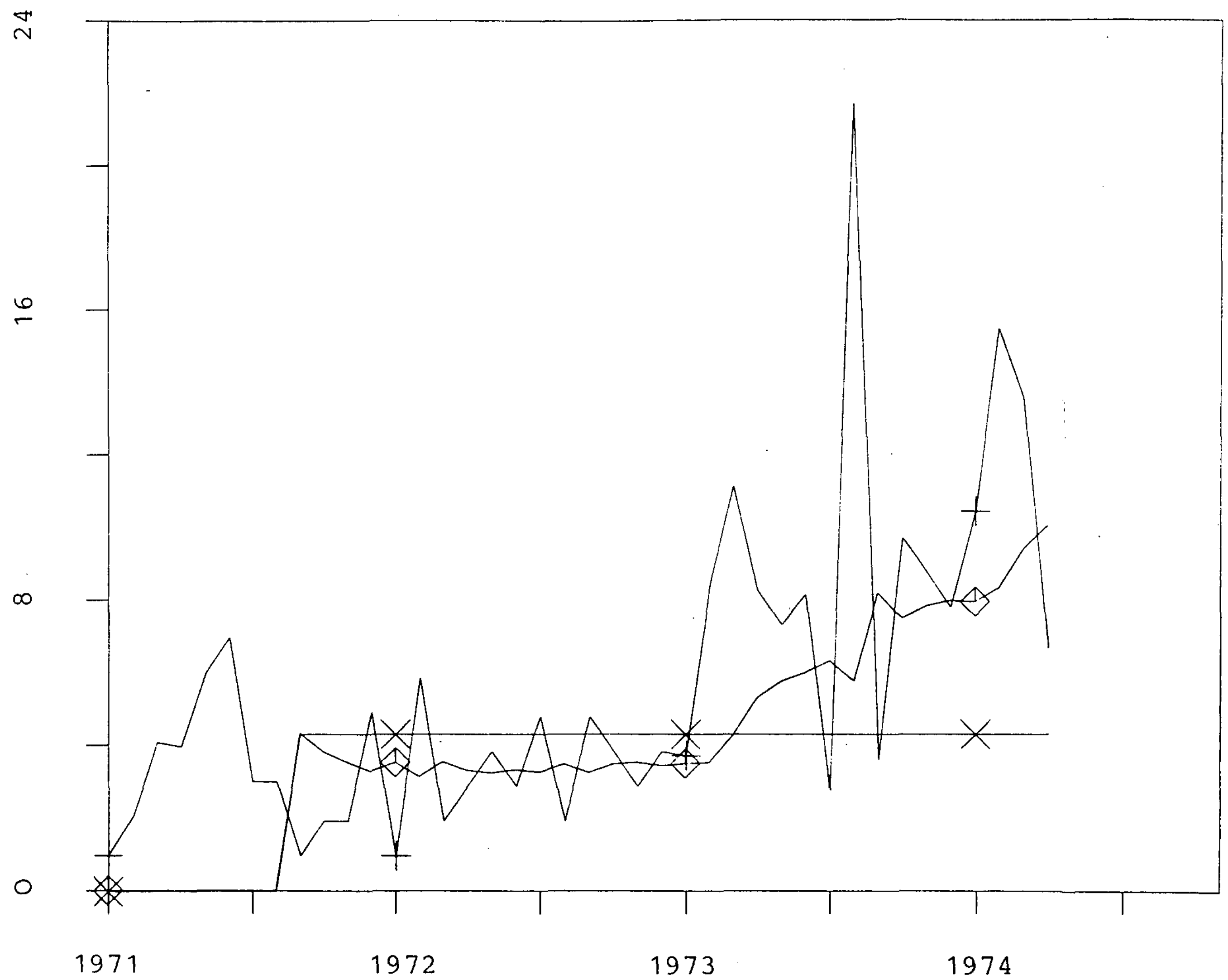
1. The Consumer Price Index CPI

Table 2 presents average values of the actual and simulated series for the different Phases and the month to month variability is depicted in Figure 1.

The results from both types of forecast is quite the same. An initial, albeit insignificant drop in the rate of CPI

* Note: The reported standard deviation of forecast errors are underestimates since we are using the estimated values from the Ψ weights and not the unobservable real ones.

FIGURE 1 : Actual and simulated annualized monthly rates of change in the consumer price index CPI



◇ updated
× deterministic
+ actual

Phase I : 1971,9-1971,11, Phase II: 1971,12-1972,12
Phase III: 1973,1-1973,6, Phase IV: 1973,7-1974,4

TABLE 2 Actual and simulated rate of change (annualized) of the Consumer Price Index CPI

(Number in brackets: standard error of forecast)

DETERMINISTIC: (ℓ-step ahead forecasts)

Period	Precontrols Average 1970,5-1971,8	Phase I	Phase II	Average I & II	Phase III	Phase IV*	All Phases Average
Actual	4.75	1.6	3.48	3.12	7.84	10.05	6.17
Simulated	-	4.33 (1.69)	4.33 (1.45)	4.33 (1.33)	4.33 (1.97)	4.33 (2.13)	4.33 (1.52)
Difference	-	-2.73	-0.85	-1.21	3.51	5.72	1.84

UPDATED: (one-step ahead forecasts)

Period	Precontrols average 1970,5-1971,8	Phase I	Phase II	Average I & II	Phase III*	Phase IV*	All Phases Average
Actual	4.75	1.6	3.48	3.12	7.84	10.05	6.17
Simulated	-	3.89 (1.38)	3.37 (0.72)	3.46 (0.64)	4.72 (1.15)	7.93 (0.95)	5.09 (1.53)
Difference	-	-2.29	0.11	-0.34	3.12	2.12	1.08

* Difference outside two standard errors of forecast.

inflation during Phase I was followed by an also insignificant impact during Phase II with a significant but positive impact during Phases III and IV. All Phases taken together give an estimate which points out towards higher rate of CPI inflation as a result of controls, but since it is not outside two standard deviations of forecast error, we cannot conclude that CPI inflation on average was really higher than what it would have been without controls.

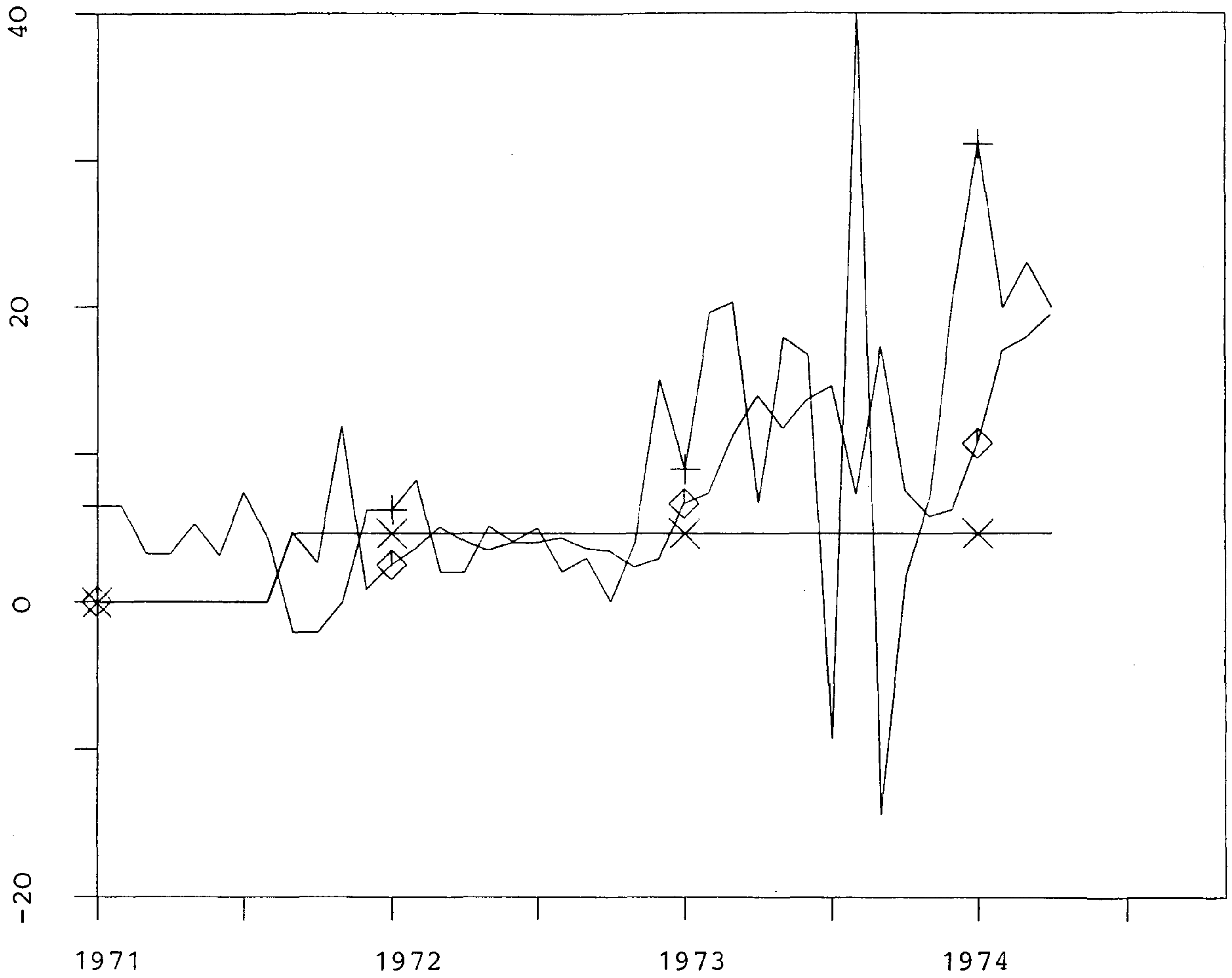
2. Wholesale Price Index (Manufacturing) WPI

As we can see from Table 3, both type of forecasts show that during Phase II the simulated values exceeded the actual by more than two standard errors, indicating that during Phase I controls seem to have been successful in bringing down the rate of WPI inflation between 6 and 8% less than what it would have been in the absence of the freeze. However, when we take into account the seemingly positive impact during Phase II the overall negative impact on WPI inflation for these two phases taken together seems negligible. Similarly to the CPI case, during Phases III and IV producer prices seem to have increased faster as a result of controls than what they would have done without them. The overall effect of all Phases was as well a positive one, although only within one standard error.

3. Average Hourly Earnings (Manufacturing) WHE

Table 4 and Figure 3 report the results for this variable. None of the differences between actual and forecast values exceeded two standard errors for any of the Phases of the ESP, although there seemed to be a sizable downward effect on the WHE rate of change during Phase I followed by quite small effects of alternated signs in the remaining different periods of controls.

FIGURE 2 : Actual and simulated annualized monthy rates of change in the wholesale Price Index (Manufacturing) WPI



◊ updated
× deterministic
+ actual

Phase I : 1971,9-1971,11, Phase II: 1971,12-1972,12
Phase III: 1973,1-1973,6, Phase IV: 1973,7-1974,4

TABLE 3 Actual and simulated rate of change (annualized) of the Wholesale Price Index WPI
(Number in brackets: standard error of forecast)

DETERMINISTIC FORECASTS: (ℓ-step ahead forecasts)

Period	Precontrols Average 1970,5-1971,8	Phase I *	Phase II	Average I & II	Phase III *	Phase IV	All Phases Average
Actual	3.48	-1.36	4.87	4.21	15.1	12.02	8.69
Simulated	-	4.70 (2.30)	4.70 (3.0)	4.70 (2.69)	4.70 (4.31)	4.70 (4.48)	4.70 (3.40)
Difference	-	-6.06	0.17	-0.49	10.4	7.32	3.99

UPDATED FORECASTS: (one-step ahead forecasts)

Period	Precontrols Average 1970,5-1971,8	Phase I *	Phase II	Average I & II	Phase III *	Phase IV	All Phases Average
Actual	3.48	-1.36	4.87	4.21	15.1	12.02	8.69
Simulated	-	6.40 (1.81)	3.41 (1.12)	3.97 (0.77)	10.7 (2.02)	12.39 (1.40)	7.86 (0.83)
Difference	-	-7.76	1.46	0.24	4.4	-0.37	0.82

FIGURE 3 : Actual and simulated annualized monthly rates of change in the manufacturing sector hourly earnings (adjusted for overtime) WHE

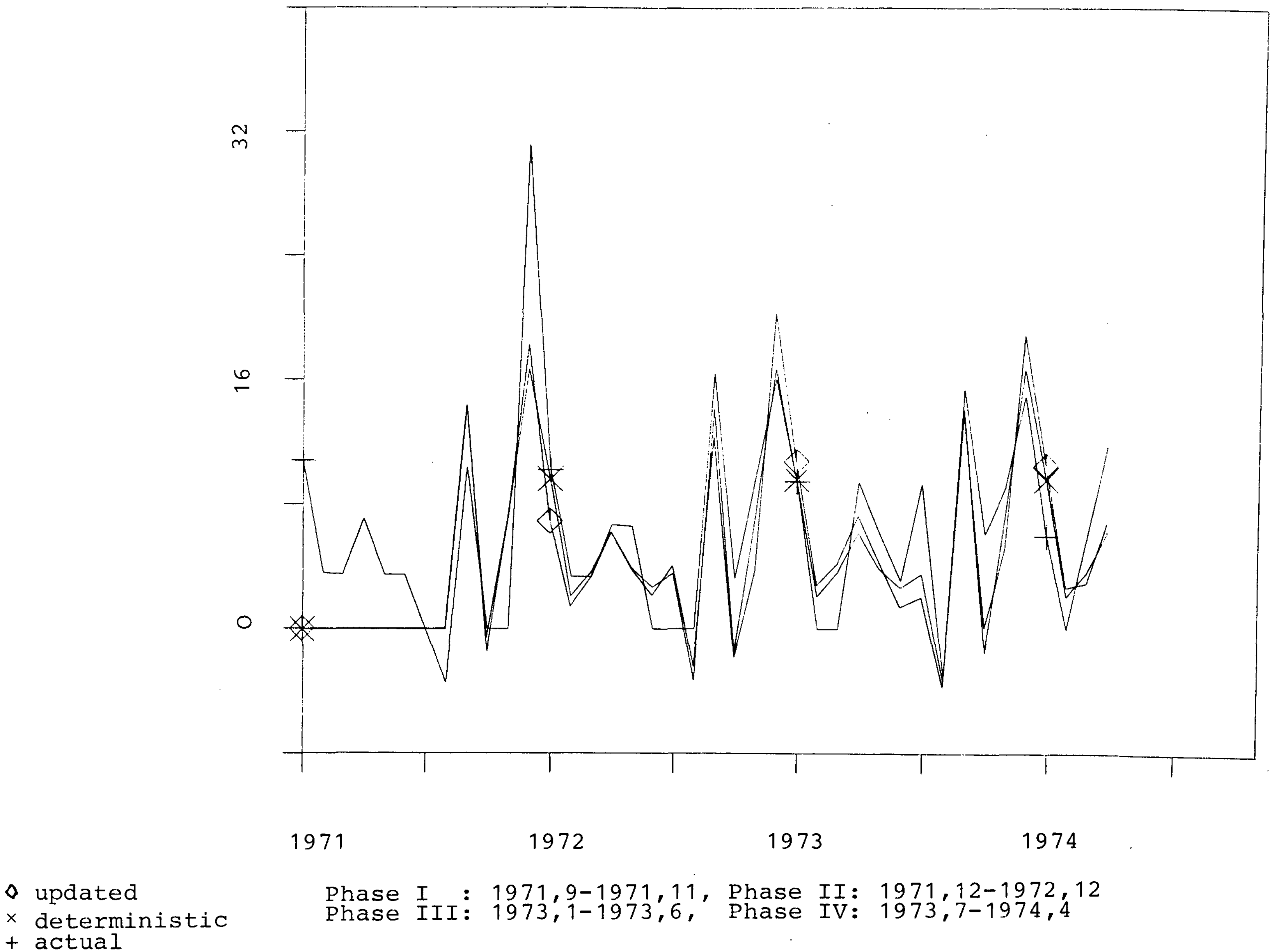


TABLE 4 Actual and simulated rate of change (annualized) of the nominal wage WHE
(Number in brackets: standard error of forecast)

DETERMINISTIC FORECASTS: (ℓ-step ahead forecasts)

Period	Precontrols Average 1970,5-1971,8	Phase I	Phase II	Average I & II	Phase III	Phase IV	All Phases Average
Actual	5.25	3.48	8.22	7.43	4.72	7.56	6.69
Simulated	-	6.72 (2.84)	6.32 (1.21)	6.39 (1.15)	4.73 (1.86)	5.87 (1.47)	5.91 (0.96)
Difference	-	-3.24	1.9	1.04	-0.01	1.69	1.05

UPDATED FORECASTS: (one-step ahead forecasts)

Period	Precontrols Average 1970,5-1971,8	Phase I	Phase II	Average I & II	Phase III	Phase IV	All Phases Average
Actual	5.25	3.48	8.22	7.43	4.72	7.56	6.96
Simulated	-	6.95 (3.05)	6.02 (1.50)	6.43 (1.34)	5.13 (2.29)	5.92 (1.84)	6.02 (0.98)
Difference	-	-3.47	2.2	1.0	-0.41	1.64	0.94

Combining the results for nominal wages WHE and prices CPI and WPI, some indication of a total average drop in the consumer real wage rate of change and a somewhat bigger decrease in the producer real wage can be detected. This result corroborates other previous evidence like that of Parkin (18) for the U.K., pointing out to a redistribution of income unfavourable to labour during control periods. One should not try to run away too easily with the conclusion that the connection between controls and the drop in the labour's share in national income is all that strong but nevertheless a type of controls like those used during the ESP may very likely have had a bigger impact on employment income than on prices of final output and consequently on profit margins.

More statistically satisfactory evidence of this type of effect is found as we shall see further below, when we look at the counterfactual forecasts of both real wage series taken on their own.

4. Manufacturing Employment EM

If one believes in long-term Phillips curves, and pushing them about, a succesful incomes policy would be judged on the basis of how effective the controls were in allowing increased employment at inflation rates below those which might have occurred without the controls. From the results in Tables 2 and 3 for the rates of CPI and WPI we have already seen that there's no conclusive evidence that these last two variables were significanly reduced and as a matter of fact they increased in the last two Phases. An increase in employment might still have been possible and in reality the actual rate of change in employment seemed to rise for the controls periods as whole as can be seen in Table 5. The initial impact of controls seemed to increase the willingness to hire workers by the unionized firms but the overall effect on employment during the ESP may also reflect a simultaneous expansionist monetary policy during specially Phase II since the rate of money growth increased substantially during this period as

FIGURE 4 : Actual and simulated annualized monthly rates of change in the manufacturing sector employment EM

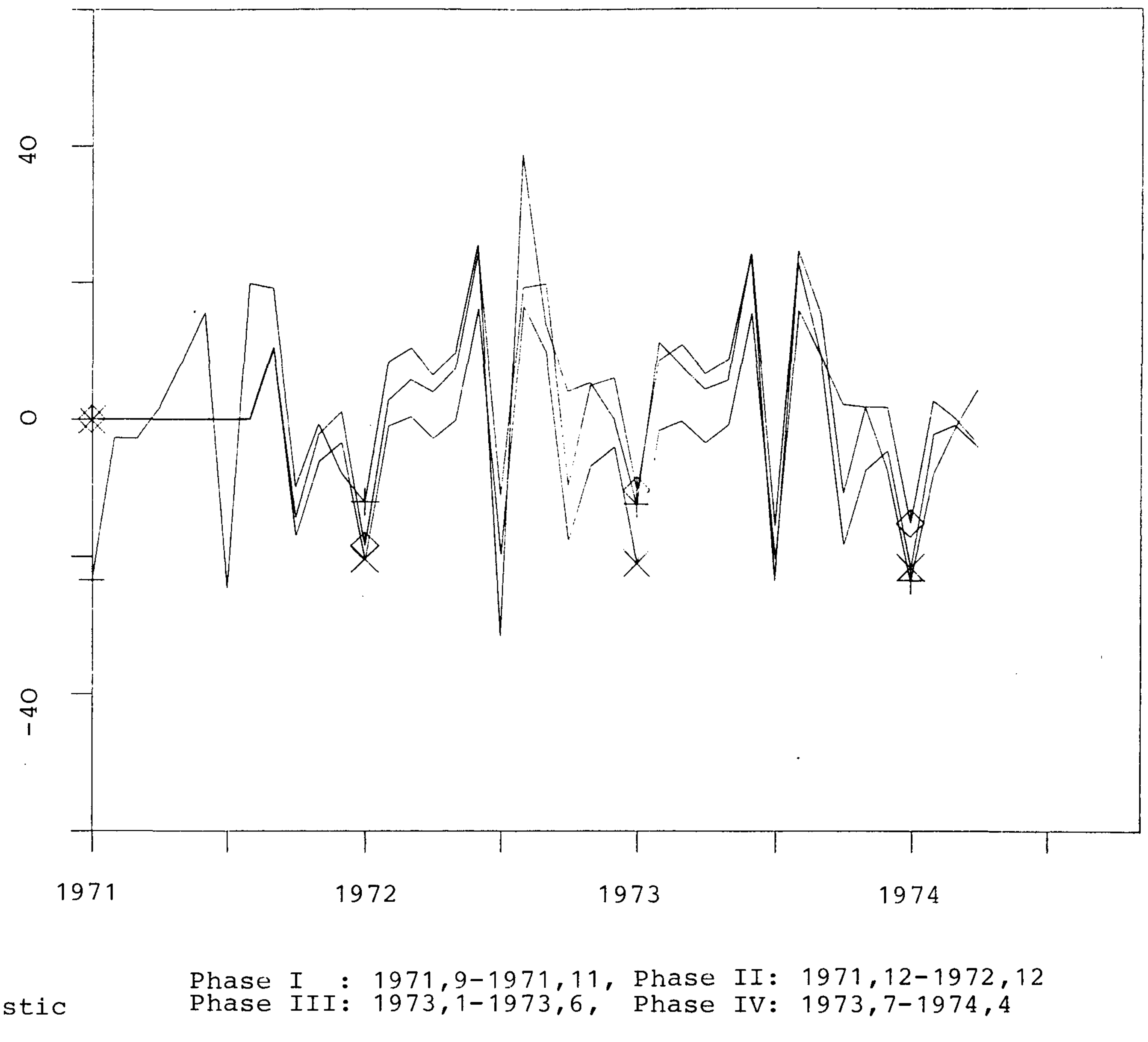


TABLE 5 Actual and simulated rate of change (annualized) of employment EM
(Number in brackets: standard error of forecast)

DETERMINISTIC FORECASTS: (ℓ-step ahead forecasts)

Period	Precontrols Average 1970,5-1971,8	Phase I	Phase II	Average I & II	Phase III	Phase IV	All Phases Average
Actual	-3.64	2.92	5.50	5.01	6.84	-2.32	3.06
Simulated	-	-4.17 (9.36)	-2.45 (12.34)	-2.77 (11.02)	1.84 (18.4)	-5.37 (22.4)	-2.71 (14.9)
Difference	-	7.09	7.95	7.78	5	3.05	5.77

UPDATED FORECASTS: (one-step ahead forecasts)

Period	Precontrols Average 1970,5-1971,8	Phase I	Phase II	Average I & II	Phase III	Phase IV	All Phases Average
Actual	-3.64	2.92	5.50	5.01	6.84	-2.32	3.06
Simulated	-	-2.04 (7.42)	4.27 (4.92)	3.08 (4.23)	8.08 (8.63)	0.05 (7.93)	3.07 (3.63)
Difference	-	4.96	1.23	1.93	-1.24	-2.37	Ø.ØØ

the reported success of the freeze during the first months of the ESP allowed more room for heating up the economy during the whole of 1972. (The rate of monetary expansion during 1972 averaged 9 % versus 2.5 % during Phase I and 6 % during 1971). When looking at the results from Table 5, we cannot conclude that controls were significantly effective (even indirectly through the just mentioned expansionary self-delusion) in increasing employment by more than it would have been without them.

It would not be, however, strictly correct to make assumptions about the effect of controls on prices and employment without some complementary assumption as to what course monetary policy would have taken without the ESP. However, I do not think that this problem could be dealt with similarly as we have already done with other variables since to use forecasts from an ARIMA model fitted to money supply before the imposition of controls in order to measure the alternative monetary policy that the Federal Reserve System would have followed in the absence of controls does seem to be neither a statistically nor a theoretically sound practice. We shall assume therefore that the course of monetary policy after controls cannot be meaningfully distinguished from its course before them.

Bearing in mind then, that controls nevertheless may have affected indirectly the course of monetary policy, one might also rationalize an increase in the rate of change in employment as partly a short term reflection of the fact that the producer real wage RWW might have been falling at a faster rate that the consumer real wage RWC and preceding it.

This is in effect what a neoclassical equilibrium monetarist theory would predict, since increases in employment through unanticipated inflation (brought about for instance by an acceleration in the money supply) require according to this short-run view of the Phillips curve that the demand for labour (identified by a producer real wage) responds always quicker than the supply of it (identified by a consumer real wage) which is equivalent to stating that a producer price index WPI changes should precede changes in the CPI.

The Tables 6 and 7 reporting the results for the producer real wage RWW and the consumer real wage RWC seem to bring some evidence on these just mentioned possible effects although the lack of parallel evidence on Table 5 about significant increases in the rate of change in employment is not necessarily at odds with the last argument due to the necessarily lagged response of employment to changes in the real wage (Fixed length of contracts, labour adjustment costs, etc...)

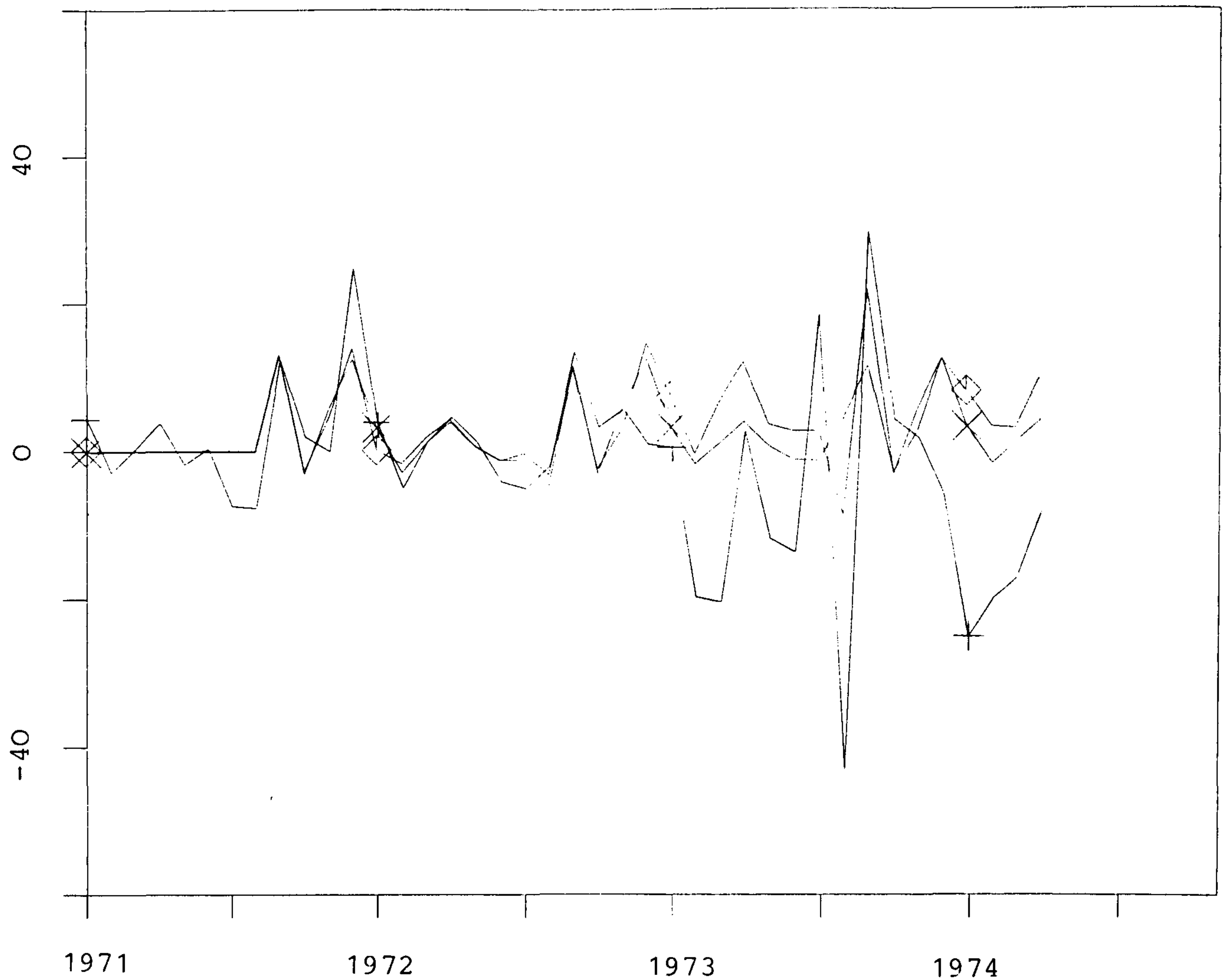
5. Producer Real Wage RWW

The estimates of Table 6, point out to a significant decrease in the producer real wage of a quite substantial magnitude in Phases III and IV and also significant drop for the whole ESP considered as a whole. An ambivalent although unsignificant effect was estimated for Phases I and II.

6. Consumer Real Wage RWC

In Table 7, one can observe that the consumer real wage appears to have increased significantly after the first two Phases, thus corroborating Gordon results (10) in which he concluded that labour might have been favoured initially by the controls and giving counterevidence to the results of Feige & Pearce (9) which stated just the opposite. However as we move on to the results for Phase III and IV we can see that by the last Phase the consumer real wage had also dropped significantly although not by as much an amount as the producer real wage RWW, which also if we remember, was also significantly reduced during Phase III.

FIGURE 5 : Actual and simulated annualized monthly rates of change in the producer real wage (manufacturing) RWW



◇ updated
× deterministic
\+ actual

Phase I : 1971,9-1971,11, Phase II: 1971,12-1972,12
Phase III: 1973,1-1973,6, Phase IV: 1973,7-1974,4

TABLE 6 Actual and simulated rate of change (annualized) of the producer real wage RWW
(Number in brackets: standard error of forecast)

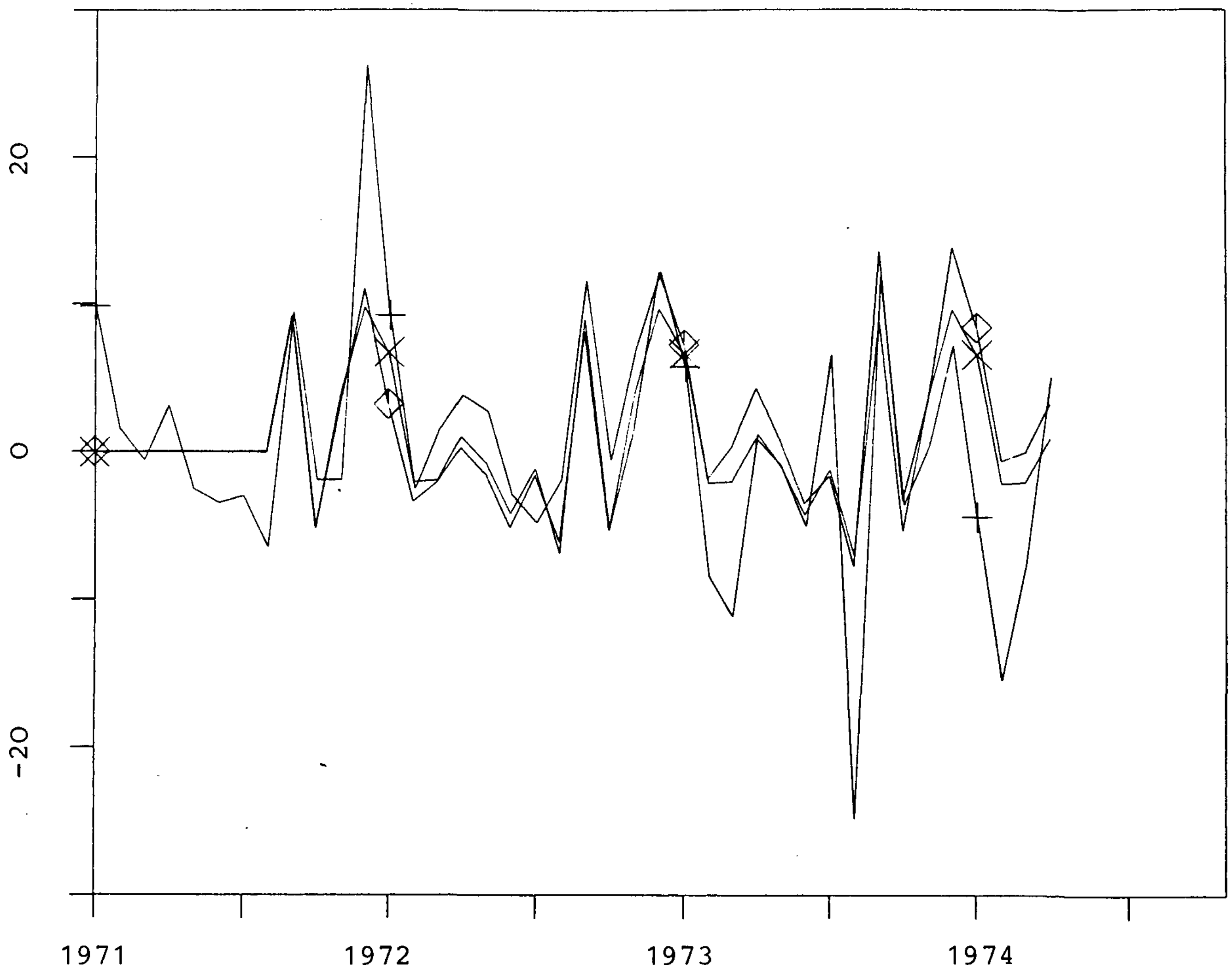
DETERMINISTIC FORECASTS: (ℓ-step ahead forecasts)

Period	Precontrols Average 1970,5-1971,8	Phase I	Phase II	Average I & II	Phase III *	Phase IV *	All Phases* Average
Actual	1.77	4.84	3.35	3.62	-10.38	-4.46	-1.52
Simulated	-	5.37 (2.59)	3.17 (1.01)	3.58 (0.94)	1.15 (2.03)	2.88 (1.36)	2.90 (0.55)
Difference	-	-0.53	0.18	0.04	-11.53	-7.34	-4.42

UPDATED FORECASTS: (one-step ahead forecasts)

Period	Precontrols Average 1970,5-1971,8	Phase I	Phase II	Average I & II	Phase III *	Phase IV *	All Phases* Average
Actual	1.77	4.84	3.35	3.62	-10.38	-4.46	-1.52
Simulated	-	4.97 (3.24)	3.16 (1.65)	3.49 (1.47)	5.40 (2.78)	5.24 (2.24)	4.39 (1.14)
Difference	-	-0.13	0.19	0.13	-15.78	-9.7	-5.91

FIGURE 6 : Actual and simulated annualized monthly rates of change in the consumer real wage (manufacturing) RWC



◇ updated
× deterministic
+ actual

Phase I : 1971,9-1971,11, Phase II: 1971,12-1972,12
Phase III: 1973,1-1973,6, Phase IV: 1973,7-1974,4

TABLE 7 Actual and simulated rate of change (annualized) of the consumer real wage RWC
(Number in brackets: standard error of forecast)

DETERMINISTIC FORECASTS: (ℓ-step ahead forecasts)

Period	Precontrols Average 1970,5-1971,8	Phase I	Phase II*	Average* I & II	Phase III	Phase IV*	All Phases Average
Actual	0.5	1.88	4.74	4.20	-3.12	-2.49	0.73
Simulated	-	2.65 (2.73)	1.43 (1.16)	1.65 (1.09)	-0.22 (1.87)	1.214 (1.15)	1.16 (0.90)
Difference	-	-0.77	3.31	2.55	-2.9	-3.70	-0.43

UPDATED FORECASTS: (one-step ahead forecasts)

Period	Precontrols Average 1970,5-1971,8	Phase I	Phase II*	Average* I & II	Phase III	Phase IV*	All Phases Average
Actual	0.5	1.88	4.74	4.20	-3.12	-2.49	0.73
Simulated	-	2.37 (3.28)	0.85 (1.53)	1.13 (1.39)	1.20 (2.37)	2.89 (1.88)	1.69 (1.01)
Difference	-	-0.49	3.89	3.07	-4.32	-5.38	-0.96

An all Phases average of the effect of controls upon the consumer real wage came out negative but not significantly so, a result which evidently takes into account the initial increase of the RWC during the first two Phases.

VII. AN INTERVENTION ANALYSIS OF THE IMPACT OF CONTROLS

The so called intervention analysis developed by Box and Tiao (3), gives also a complementary test to pick up some possible effect of incomes policy on the variables we have just been considering.

Intervention models are basically transfer function models in which the input variables are dummy variables normally called "Pulses" or "steps". The first one is used to model a temporary change in the level or slope of a time series (e.g. a strike) and the last one is used to test for a permanent change in the level of a time series (like in our case due to the imposition of controls) and also in its slope.

Like normal dummies they take a value of "1" for the period of incomes policy and "0" elsewhere, but the theoretical justification of intervention models is somewhat different as in the case of normal intercept or slope dummies combined with standard econometric models. To illustrate the way intervention analysis works, let us break down a series Y_t into two components D_t and N_t

$$y_t = D_t + N_t \qquad (11)$$

where D_t = part of y_t which can be determined exactly in terms of any input X_t.

N_t = "noise" or error term which cannot be explained in terms of X_t. It picks up all the "omitted" X-variables.

A general way to represent the relationship between D_t and X_t is with a linear dynamic relationship of the form

$$D_t = \frac{\omega(B)}{\delta(B)} X_{t-b} = V(B) X_t \qquad (12)$$

where the transfer tunction $V(B) = \frac{\omega(B)}{\delta(B)}B^b$ consists of a moving average $\omega(B)$ operator, an autoregressive operator $\delta(B)$ and a pure delay parameter b representing the number of periods before the change in the input X_t begins to have an effect on Y_t.

In general, the noise will be non-stationary and may be represented normally by an ARIMA (p,d,q)

$$\Delta^d N = \theta(B)\, \emptyset\, (B)^{-1} \varepsilon_t \qquad (13)$$

which in combination with the model for D_t yields a transfer function-noise model.

Now if we want to investigate the effect of incomes policy by using an intervention "step" variable ξ_t as input we can postulate initially a transfer function of the form

$$y_t = \omega(B)\delta^{-1}(B)\xi_{t-b} \qquad (14)$$

whose parameters can be estimated by the methods described in Box and Jenkins (   ). However since we cannot use prewhitening to identify the structure of model (14) like in the case of normal transfer function models, the problem has to be tackled by introducing the noise structure N_t into the intervention model. To do so, we assume a process generating a series without any abnormal event being yet considered. Referring back to the estimated model for the rate of change in employment EM described in section V., we can represent it as

$$\Delta\Delta^{12} y_t = (1-\theta_1 B)(1-\theta_{12}B^{12})\varepsilon_t \qquad (15)$$

where the model fitted to the series y_t = rate of change in employment, does not include the incomes policy periods. It follows that if the intervention mechanism was (14) we can write a model

$$y_t = \omega(B)\delta^{-1}(B)\xi_{t-b}+N_t \qquad (16)$$

where the noise term N_t describes how the series y_t is generated in the absence of incomes policy.

If the parameters in the polynomials $\omega(B)$ and $\delta(B)$ are zero we can then obtain a first guess of the structure of the noise N_t

$$\Delta\Delta^{12}N_t = (1-\theta_1 B)(1-\theta_{12}B^{12})\varepsilon_t \qquad (17)$$

which in combination with equation (14) gives us an equation such as

$$\Delta\Delta^{12}y_t = \omega(B)\delta^{-1}(B)\Delta\Delta^{12}\xi_{t-b}+(1-\theta_1 B)(1-\theta_{12}B^{12})\varepsilon_t$$

or alternatively

$$y_t = \omega(B)\delta^{-1}(B)\xi_{t-b}+\frac{(1-\theta_1 B)(1-\theta_{12}B^{12})\varepsilon_t}{\Delta\Delta^{12}} \qquad (18)$$

which yields maximum likelihood estimates of the parameters in $\omega(B)$, $\delta(B)$ and the parameters θ, by minimising the sum of squared residuals $\hat{\varepsilon}_t$ by nonlinear methods.

Models like (18), assuming that only $\omega_o \neq o$ in the polynomial $\omega(B)$, not including parameters of $\delta(B)$ and with a parameter b=o would be thus equivalent to test for the hypothesis that controls bring about an upward or downward shift in the level of the variables we are studying, but leaves the underlying process generating them unchanged in this sense a model such as

$$y_t = \omega_o \xi_t + \frac{(1-\theta_1 B)(1-\theta_{12} B^{12})}{\Delta\Delta^{12}} \varepsilon_t \qquad (19)$$

where $\xi_t = 1$ in all the months of the controls program, tests for some significant average effect of incomes policy on the rate of change of employment without taking into account the time pattern of adjustment to a new level which would be allowed for instance by a model such as

$$y_t = \omega_o (1-\delta B)^{-1} \xi_t + \frac{(1-\theta_1 B)(1-\theta_{12} B^{12}) \varepsilon_t}{\Delta\Delta^{12}} \qquad (20$$

Models like (19) for the rate of change in employment, were estimated for all the other variables, CPI, WPI, WHE, RWW and RWC with the noise model N_t varying across them as can be immediately seen from the fitted ARIMA models in Table 1.

According to whether the separate effects of the four phases of the controls, the first two lumped together or the overall effect of controls were taken into account, the results are respectively reported in Tables 8, 9 and 10.

A comforting outcome is that the results of the intervention analysis are very much in line with the results from the counterfactual simulations already reported in section VI. Magnitude, sign and significance of coefficients are a confirmation of the results we have already seen and commented upon in section VI; with the most striking result again, the negative impact on both real wages but of a bigger magnitude and unambiguous in the case of the producer real wage RWW.

TABLE 8 INTERVENTION ANALYSIS (Different Phases separately)
Maximum likelihood estimates and associated standard errors

	Phase I $\hat{\omega}_{o1}$	Phase II $\hat{\omega}_{o2}$	Phase III $\hat{\omega}_{o3}$	Phase IV $\hat{\omega}_{o4}$	$\hat{\theta}_1$	$\hat{\theta}_{12}$
CPI	-2.85 (1.69)	-0.97 (1.36)	3.25 (1.81)	5.27* (1.88)	0.88* (0.03)	
WPI	-5.71 (3.3)	0.73 (3.18)	10.04* (4.22)	5.70 (4.62)	0.81* (0.040)	
WHE	-2.88 (2.44)	1.68 (1.21)	0.56 (1.76)	1.62 (1.43)	0.20* (0.066)	0.68* (0.056)
EM	5.88 (8.91)	6.36 (10.1)	6.19 (13.67)	1.28 (15.9)	0.71* (0.04)	0.82* (0.04)
RWW	-1.61 (3.04)	0.20 (1.39)	-9.95* (2.08)	-9.71* (1.62)	0.27* (0.06)	0.79* (0.05)
RWC	-0.21 (3.6)	2.94* (1.29)	-2.26 (1.87)	-3.93* (1.52)	0.25* (0.06)	0.78* (0.04)

* = significant at 0.05 level

TABLE 9 INTERVENTION ANALYSIS (Phase I & Phase II combined)
Maximum likelihood estimates and associated standard errors

	Phase I & II $\hat{\omega}_{01}$	Phase III $\hat{\omega}_{02}$	Phase IV $\hat{\omega}_{03}$	$\hat{\theta}_1$	$\hat{\theta}_{12}$
CPI	-1.59 (1.25)	2.83 (1.77)	4.85* (1.84)	0.88* (0.02)	
WPI	-2.28 (2.84)	7.29 (4.08)	2.84 (4.52)	0.81* (0.04)	
WHE	0.93 (1.11)	0.39 (1.76)	1.63 (1.44)	0.20* (0.06)	0.67* (0.09)
EM	6.06 (8.3)	5.85 (12.01)	0.93 (14.5)	0.71* (0.04)	0.82* (0.04)
RWW	-0.08 (1.28)	-10.0* (2.07)	-9.2* (1.62)	0.27* (0.06)	0.79* (0.05)
RWC	1.96 (1.15)	-2.87 (1.84)	-4.51* (1.46)	0.24* (0.06)	0.76* (0.04)

* = significant at 0.05 level

TABLE 10 INTERVENTION ANALYSIS (All Phases combined)

Maximum likelihood estimates and associated standard errors

	All Phases $\hat{\omega}_{01}$	$\hat{\theta}_1$	$\hat{\theta}_{12}$
CPI	-1.49 (1.55)	0.81* (0.07)	
WPI	-2.56 (3.07)	0.78* (0.04)	
WHE	1.02 (0.95)	0.20* (0.06)	0.67* (0.05)
EM	6.43 (8.23)	0.71* (0.04)	0.82* (0.04)
RWW	-4.76* (1.29)	0.14* (0.06)	0.78* (0.05)
RWC	-0.78 (1.06)	0.16* (0.06)	0.76* (0.04)

* = significant at 0.05 level

VIII. CAUSALITY TESTS

Referring back to the problem already mentioned in section II and III about the observed rates of price inflation not reflecting the real ones, because of changes in the quality of products, it is an interesting exercise in itself to try to find some empirical evidence of the effect of this type of phenomenon on employment decisions. Firms basing their decisions to hire or lay off workers upon the producer real wage RWW would not do so by taking into account the reported rate of inflation in WPI but rather they will consider prices in their output which already allow for the change in quality they may have caused so as to restore their desired profit margins (only of course when the ceiling on the prices of the output they sell is binding enough to give them an incentive to do so).

The message in money wages data may as well be affected by controls, since firms and employees will find it relatively easy to evade wage controls as long as they find it beneficial to do so. In competitive labour markets employers will be interested in dodging controls by means such as labour upgrading or overreported number of hours worked if they think they will lose valuable employees otherwise. In the unionized U.S. manufacturing sector however, and during the first two Phases at least, firms seem to have been able to hire the number of employees they wanted under the enforced wage rate and therefore the incentive to evade controls was not so strong.

The last argument points towards a misreported real wage coming more from prices than from wages.

We have already seen in section II that firms seemingly started to restore quality during Phase III and Phase IV. The overall effect on the reported rate of inflation is not clear-cut, although it might have been also misreported for the ESP taken as a whole . An attempt to bring out some evidence this respect can be carried out by means of Granger causality tests (12) of the Pierce-Haugh (20) variety between real wages and employment.

In short, the basic idea can be put down as follows:
Let us imagine that the process generating the rate of change in employment can be modelled like

$$y_t = \phi_2^{-1}(B)\,\theta_2(B)\,\varepsilon_{2t}$$

The residuals ε_{2t} can be viewed as that part of y_t which cannot be predicted from its own past history. Therefore if one were to improve the forecasts of y_t using extra information, like for instance, a real wage X_t modelled by

$$X_t = \phi_1^{-1}(B)\,\theta_1(B)\,\varepsilon_{1t}$$

The shocks or innovations ε_{1t} should be correlated with ε_{2t} (Reason for not simply cross-correlating the prewhitened y_t series, ε_{2t} with the X_t series is that the resulting cross-correlation function may be misleading due to serial correlation in the X_t series (see Bartlett (23)).

When a series such as X_t provides additional information about future values of y_t or in other words reduces the expected mean square forecast error below that of a model of y_t based only on its own past history, we can say that X_t is a "leading indicator" of y_t or that satisfies the condition stated by Granger (11) necessary to say that X causes y. (For a more detailed description see Pierce and Haugh (20)).

The idea of Granger causality however is based on the use of an information set which may be bigger than just two variables X and y. Therefore discussing causality in the way we have done so far, would be equivalent to consider information sets with only two elements; that is to consider causal relationships pair-wise amongst a selected group of variables. With this proviso in mind one can then proceed further and test for the existence or not of causality by cross-correlating the innovations ε_t of the processes generating the variables in our information set.

The cross-correlogram between both prewhitened series can be written as

$$\rho_{\varepsilon 1 \varepsilon 2, k} = E(\varepsilon_{1t} \varepsilon_{2t+k}) / \sigma_{\varepsilon 1} \cdot \sigma_{\varepsilon 2}$$

where k = positive, negative or zero lag.

As shown by Haugh (13), if $\rho_{\varepsilon 1 \varepsilon 2, k} \neq o$ for some $k > o$ one can say that "X causes y", on the other hand if $\rho_{\varepsilon 1 \varepsilon 2, k} \neq o$ for some $k < o$ then "y causes X". If for both $k < o$ and $k > o$ $\rho_{\varepsilon 1} \rho_{\varepsilon 2, k} \neq o$ there's feedback, and finally when $\rho_{\varepsilon 1} \rho_{\varepsilon 2, o} \neq o$ the direction of causality is indeterminate and nothing can be concluded. Only when $\rho_{\varepsilon 1 \varepsilon 2, k} = o$ for all k can one say that the series y_t and X_t, in our case employment and real wages, are strictly independent. However, as pointed out by Sims (25) these may be a tendency for the crosscorrelations to be biased towards zero due to specification error.

Since the population cross-correlogram $\rho_{\varepsilon 1 \varepsilon 2,k}$ is unobservable we estimate it with the sample residual cross-correlogram

$$\hat{r}_{\hat{\varepsilon}1\hat{\varepsilon}2,k} = \frac{\Sigma \hat{\varepsilon}_{1t}\ \hat{\varepsilon}_{2t+k}}{\left[(\Sigma \hat{\varepsilon}^{2}_{1t}) \cdot (\Sigma \hat{\varepsilon}_{2t})\right]^{1/2}}$$

which is in turn an estimate of the sample white noise cross-correlation $r_{\varepsilon 1 \varepsilon 2,k}$.

Two basic tests can be employed then to detect causality. The first is basically to check whether the individual $\hat{r}$'s exceed the value $2\ T^{-1/2}$ since under the null hypothesis of independence between both series

$$\hat{r}_{\hat{\varepsilon}1\hat{\varepsilon}2,k} \sim N(0,T^{-1})$$

where T = number of observations.

Therefore it is straightforward to say that we are approximately 95% certain of rejecting the null hypothesis when it is false when we obtain some

$$\hat{r}_{\hat{\varepsilon}1\hat{\varepsilon}2,k} > 2T^{-1/2}$$

A second, "portmanteau" test, follows immediately by adding up the squared values of the correlogram, since the assumption of white noise normality yields

$$S = T \sum_{k=-p}^{p} \hat{r}^{2}_{\hat{\varepsilon}1\hat{\varepsilon}2}(k) \sim \chi^{2}(2p+1)$$

which can be compared with the relevant value of the χ^2 tables to accept or reject the hypothesis of independence.

Correcting for small sample bias and taking into account the sensitivity of the test to the choosing of lags P, Layton (15) suggests a scheme of progressive testing such as

a) X causes y if any one of

$$S(p) = T^2 \sum_{k=1}^{p} (T-|k|)^{-1} \hat{r}^2(k) \sim \chi^2(p)$$

is significant

b) y causes X if any one of

$$S(p) = T^2 \sum_{k=-1}^{-p} (T-|k|)^{-1} \hat{r}^2(k)^2 \sim \chi^2(p)$$

is significant

c) X and y are independent if none of

$$S(p) = T^2 \sum_{k=-p}^{p} (T-|k|)^{-1} \hat{r}^2(k) \sim \chi^2(2p+1)$$

is significant.

If unidirectional causality is detected, then one should confirm it by fitting a transfer function-noise model (see Haugh (13), Box-Jenkins (2)) like the one already mentioned in section VII

$$y_t = \omega(B)\delta(B)^{-1} X_{t-b} + \theta(B)\phi^{-1}(B)\varepsilon_{3t}$$

And the residuals $\hat{\varepsilon}_{3t}$ obtained from this model should be cross-correlated with $\hat{\varepsilon}_{2t}$ at positive and negative lags to confirm unidirectional causality using a similar χ^2 test as the one explained before.

As we see in Tables 11 and 12 to use the second step does not appear to promise a substantial pay-off since evidence on clear cut unidirectional causality between both real wages RWC and RWW and employment is non-existent from models fitted to the period prior to the imposition of controls 1953,7 - 1971,8.

TABLE 11 Cross-Correlation between residuals from RWC and EM estimated models 1953,7 - 1971,8 : Precontrols ($2\hat{\sigma} = 0.139$)

k :	0	1	2	3	4	5	6	7	8	9	10	11	12	13	14	15	16
$\hat{r}_k$:	0.45*	0.14*	0.09	-0.04	0.04	-0.02	0.01	0.07	0.02	0.03	-0.04	-0.03	-0.04	-0.07	0.07	-0.07	0.07
k :	17	18	19	20	21	22	23	24	25	26	27	28	29	30			
$\hat{r}_k$:	0	0.10	0.05	-0.07	0	-0.06	0.03	-0.01	0	0.01	0	0	-0.09	-0.01			
k :		-1	-2	-3	-4	-5	-6	-7	-8	-9	-10	-11	-12	-13	-14	-15	-16
$\hat{r}_k$:		-0.14*	-0.09	-0.06	-0.11	-0.01	-0.01	0.07	0.03	0.03	0.03	0	-0.12	-0.01	0	0	0.06
k :		-17	-18	-19	-20	-21	-22	-23	-24	-25	-26	-27	-28	-29	-30		
$\hat{r}_k$:		0	0.06	-0.13+	-0.07	-0.06	-0.10	-0.02	-0.08	0.06	0.04	0.08	0.07	0	-0.02		

TABLE 12 Cross-Correlation between residuals from RWW and Em estimated models 1953,7 - 1971,8 : Precontrols ($2\hat{\sigma} = 0.139$)

k	0	1	2	3	4	5	6	7	8	9	10	11	12	13	14	15	16
$\hat{r}_k$	0.37*	0.03	0.02	−0.03	−0.07	−0.06	0	0.03	0	0	−0.03	−0.04	−0.06	−0.13+	0.02	−0.07	0.04
k	17	18	19	20	21	22	23	24	25	26	27	28	29	30			
$\hat{r}_k$	0.08	0.1	0.07	−0.04	−0.02	−0.05	0.02	−0.06	−0.01	−0.01	0.03	0.05	−0.03	0.01			
k		−1	−2	−3	−4	−5	−6	−7	−8	−9	−10	−11	−12	−13	−14	−15	−16
$\hat{r}_k$		−0.12	−0.10	−0.03	−0.07	0.02	0.03	−0.05	0.04	0.08	0	−0.08	0.05	−0.03	0.01	0.11	0.06
k	−17	−18	−19	−20	−21	−22	−23	−24	−25	−26	−27	−28	−29	−30			
$\hat{r}_k$	0.02	0	−0.03	0	0	−0.07	0	−0.09	0.04	0.01	0.02	0.05	−0.01	−0.04			

The contemporaneous cross-correlation is significant in both Tables, but nothing can be inferred from that fact as regards causality. A weak evidence of feedback is detected between the consumer real wage RWC and employment EM. However, in the crosscorrelogram between the prewhitened producer real wage and employment an almost significant at 0.05 level cross-correlation is found at lag 13 with all other cross-correlations well insignificant although there's still a sizable cross-correlation in Table 12 at lag K = -1.

The χ^2 tests corresponding to Tables 11 and 12 can be summarized as follows

1. <u>RWC and EM</u>

S(61) = 94.32 for k = o,±1,±2, ...±3o

S(6o) = 51.52 for k = ±1,±2, ...±3o

where χ^2(6o) = 79.1 at o.o5 level. When excluding the significant contemporaneous crosscorrelation, the "portmanteau"test accepts the null hypothesis of independence between both series but strictly speaking we have to include k = o to conclude that both series are independent.
Of the progressive χ^2 tests a weak evidence of feedback is picked up since only

S(1) = 4.o9 for k = +1

and S(1) = 4.o7 for k = -1

are greater than χ^2(1) = 3.84 at o.o5 level.

2. <u>RWW and EM</u>

Ever the more general "portmanteau"test S(61) = 69.48 cannot reject independence although we have already seen that significant instantaneous crosscorrelation also exists. Of the progressive χ^2 tests only S(1) = 3.84 for k = 1 allows us to say something about a weak keynesian as opposed to neoclassical causality, running from employment to real wages (with a negative sign if we remember the cross-correlogram in Table 12.

With the lack of substancial evidence about a clear line of causality between both definitions of the real wage and employment it does not seem very promising to analyze separately how controls might have affected something whose existence proves to be rather elusive even without them. Nevertheless ARIMA models for the same three series EM, RWC, RWW were refitted up to the end of controls in April 1974. The results are reported in Tables 13 and 14 where one can see that apart from significant instantaneous causality, the weak evidence of causal relationships in Tables 11 and 12 (which omit the control's period) is gone when we take the observations for the control's time span into account. Therefore some diluted evidence exists that during the ESP, the reported real wages did not reflect the "actual" real wages which, producers and consumers alike would have included in their information sets.

TABLE 13 Cross-Correlation beween residuals from RWC and EM estimated models 1953,7 - 1974,4 : with Controls ($2\hat{\sigma}$ = 0.130)

k	0	1	2	3	4	5	6	7	8	9	10	11	12	13	14	15	16
$\hat{r}_k$	0.43*	0.12	0.08	0.04	-0.04	-0.02	0.01	0.08	0.05	0.04	-0.05	-0.02	-0.02	0	0.03	-0.07	0.07
k	17	18	19	20	21	22	23	24	25	26	27	28	29	30			
$\hat{r}_k$	-0.01	0.09	0.02	-0.06	0.01	-0.08	0	-0.03	0	0.02	-0.05	-0.01	-0.08	-0.02			
k		-1	-2	-3	-4	-5	-6	-7	-8	-9	-10	-11	-12	-13	-14	-15	-16
$\hat{r}_k$		-0.08	-0.08	-0.04	-0.09	0	0.02	-0.09	0.03	0.03	0.05	0	-0.11	0.01	0.01	-0.03	0.03
k		-17	-18	-19	-20	-21	-22	-23	-24	-25	-26	-27	-28	-29	-30		
$\hat{r}_k$		-0.01	0.05	-0.05	-0.04	-0.05	-0.09	-0.02	-0.11	0.04	0.02	0.07	0.04	0	-0.03		

TABLE 14 Cross-Correlation between residuals from RWW and EM estimated models
1953,7 - 1974,4 : with controls ($2\hat{\sigma} = 0.130$)

k	0	1	2	3	4	5	6	7	8	9	10	11	12	13	14	15	16
$\hat{r}_k$	0.33*	0.02	0.02	-0.02	-0.05	0	0	0.04	0.04	0.02	-0.04	0.02	-0.07	0	-0.01	-0.07	0.03
k	17	18	19	20	21	22	23	24	25	26	27	28	29	30			
$\hat{r}_k$	0.04	0.07	0	-0.02	0	-0.04	0	-0.10	-0.01	0	-0.06	0.04	0.01	-0.02			
k		-1	-2	-3	-4	-5	-6	-7	-8	-9	-10	-11	-12	-13	-14	-15	-16
$\hat{r}_k$		-0.02	-0.08	0	-0.05	0.02	0.05	-0.04	0.02	0.07	0.02	0	-0.03	0.04	-0.01	-0.03	0.04
k		-18	-19	-20	-21	-22	-23	-24	-25	-26	-27	-28	-29	-30			
$\hat{r}_k$		0.01	-0.02	0	0	-0.05	0.01	-0.08	0.02	0	0	0.01	0	-0.02			

IX. Summary

The empirical analysis of incomes policy has normally been carried out on the basis of wage and price equations which are given a structural interpretation. This implies that all the other variables in this system apart from the endogenous wages, prices and also at times price expectations, are to be considered as exogenous. It is argued in the text that it may not be legitimate to consider as exogenous, variables such as unemployment, productivity and prices of imports. The imposition of controls may affect directly the path of these assumed exogenous variables during the simulation performed to quantify the impact of controls on the endogenous variables. This simulation is traditionally executed by the use of reduced forms which of course are obtained by previously imposing a certain "structure" on a set of variables of interest. The use of ARIMA models is defended as a short cut to the problem of imposing too much "a priori" structure. As proved in the text along the lines of Zellner and Palm (22) each of the stochastic variables belonging to a dynamic simultaneous system may be given an ARIMA form under very general conditions.

ARIMA models can therefore be used as forecasting tools which are theoretically compatible with the "true" structure of the model. The results, using ARIMA-generated counterfactual forecasts, give some further evidence that controls were not succussful in fighting down wage or price inflation (apart from some initial success in Phase I). Unemployment did not seem either to decrease significantly as a result of imposing controls.

An indirect influence of wage and price controls on the evolution of real wages through an expansionist monetary policy during Phase II may have taken place. Policy makers, apparently mistook the first information to arrive on wages and prices during Phase I, for an "all-clear" signal to stimulate

demand. The evidence on Phases III and IV is consistent with what a neoclassical equilibrum theory would predict about the timing in the response of the producers and the consumer real wages to unanticipated monetary growth.

The results in both sign and nagnitude using counterfactual simulations based on the estimated ARIMA models were confirmed to a very high degree when using the "Intervention Analysis" technique developed by Box and Tiao (3).

The paper concludes with some causality tests between both definitions of the real wage and employment. The tests are performed using first data up to the imposition of controls and secondly using as well the observations corresponding to the ESP period. It is shown that the weak evidence of causality existing prior to the enforcement of controls, disappears when the observations belonging to the controls period are also used. It is argued that this may give some indication of an error in variables phenomenon due to the false message contained in the published price data, since the latter did not reflect the drop in product quality which took place, specially during the duration of Phase II.

<u>Statistical Appendix</u>

The variables used are annualized monthly rates of change using as a basis the following seasonally unadjusted series

1. Consumer Price Index: (Source: Business Statistics)
2. Wholesale Price Index: (Source: Business Statistics)
3. Manufacturing Employment: (Source: U.S. Dept. of Labour Business Statistics)
4. Average Hourly Earnings: (Source: Business Statistics) This last series is adjusted for overtime.

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