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THE 1971–1974 CONTROLS PROGRAM AND THE PRICE LEVEL

An Econometric Post-Mortem*

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This paper provides new empirical evidence on the effects of the Nixon wage–price controls on the price level. The major new wrinkle is that the controls are treated as a quantitative (rather than just a qualitative) phenomenon through the use of a specially-constructed series indicating the fraction of the economy that was controlled. According to the estimates, by February 1974 controls had lowered the non-food non-energy price level by 3–4 percent. After that point, and especially after controls ended in April 1974, a period of rapid 'catch up' inflation eroded the gains that had been achieved, leaving the price level from zero to 2 percent below what it would have been in the absence of controls. The dismantling of controls can thus account for most the the burst of 'double digit' inflation in non-food and non-energy prices during 1974.

1. Introduction

For better or for worse, controls over wages and prices have become a standard component of the macroeconomic policymaker's arsenal. For this reason, it is useful to examine with some care the only experience the United States has ever had with mandatory wage–price controls in time of peace. By how much did the controls program initiated by the Nixon administration in August 1971 restrain inflation? How severe was the extra 'catch up' inflation that occurred when controls were lifted? To what extent does the removal of controls account for the remarkable burst of 'double digit' inflation that we experienced in 1974? These are the questions addressed in this paper.

Sections 2, 3, and 4 are brief sections reviewing, respectively, what theory suggests controls might do to the path of the price level, the major institutional features of the 1971–1974 controls, and previous attempts to measure the effects of these controls. Section 5, the heart of the paper, describes our own model of the effects of controls and discusses the estimates. Section 6 is a brief conclusion.

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2. Controls and the price level

A useful benchmark for what we expect controls to do to the path of the price level is provided by the standard supply-and-demand diagram, in which temporary, binding controls lower prices in the short run, but result in excess demand which drives prices back to their initial levels once controls are lifted. The suggestion, once we aggregate to the macro level, is that controls might influence the price level in the manner indicated in fig. 1.

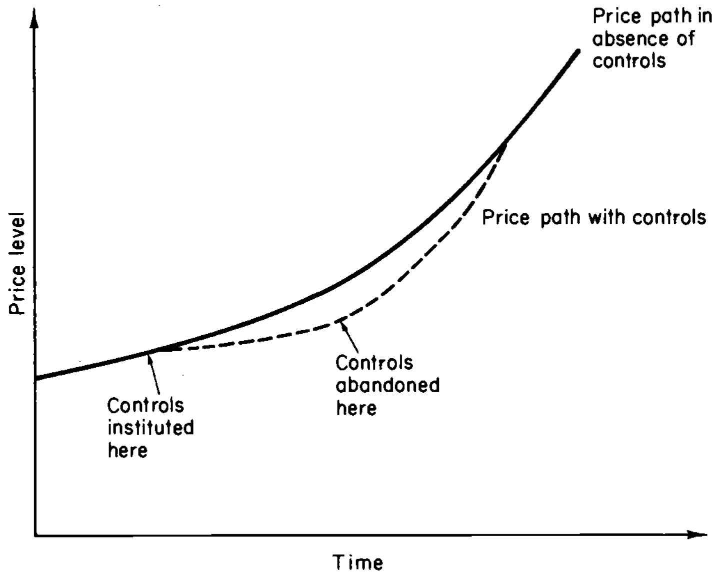

Fig. 1

Unless the controls are an exercise in futility, we certainly expect them to reduce the rate of inflation — at least as measured in the official indexes[1] — while they are in effect. How much, of course, is an empirical question. What should happen after controls are lifted is much more controversial. While simplistic use of the micro analogy suggests precisely 100% catch up, as in fig. 1, arguments can be made why the catch up might be either greater or less than 100%.

Proponents of controls point out that a controls program will also hold down marginal costs as well as product prices so that, when controls are lifted, there may not be any excess demand at the controlled price. In the

[1]If controls cause shortages which are partially alleviated by disguised price hikes (e.g., quality deterioration, ending previously-offered discounts, queueing, etc.), official price indices may understate the inflation in the 'shadow prices' [see Oi (1976)] or in a more accurate index of actual transactions prices [see Darby (1976)].

particular case of the 1971–1974 controls, the weight of empirical evidence, however, seems to be that controls squeezed profit margins, that is, were more effective on prices than on wages.[2] In such a case, marginal costs would not have been held down by as much as prices, so some catch up would be expected.[3]

Furthermore, it is virtually certain that any real-world controls program will distort relative prices, and it is abundantly clear that the 1971–1974 controls did so. To the extent that prices are rigid downward, equilibrium *relative* prices can only be re-established through rising *nominal* prices.

A different argument for complete catch up is based on the long-run neutrality of money. If the controls program does not imply a change in the long-run path of the money supply (or of fiscal policy), it cannot affect the long-run price level because (a) there is no mechanism by which temporary controls can permanently affect velocity, and (b) the natural rate hypothesis implies that real output will be at its 'full-employment' level in the long run regardless of whether or not there ever was a controls program. Supporters of controls counter that the money supply path need not be fixed. Specifically, by permitting the price level path to be deflected downward for a while, controls allow the economy to get by with less money. If the money supply path is permanently lowered, then the price level path should also be permanently lowered, in contradiction to fig. 1. In fact, however, there is no evidence that monetary growth was throttled back after August 1971.

In the arguments reviewed so far, potential *GNP* was assumed to be unaffected by controls. But this might not be so. It has been alleged (though not, to our knowledge, demonstrated empirically) that the controls discouraged business investment. The common belief that the 1971–1974 controls program squeezed profit margins by holding down prices more than wages has been used to support this allegation. If true, this is a reason why the catch up could conceivably exceed 100 percent, for in that case controls would have reduced the natural rate of real output.

Finally, the most important intellectual justification for instituting controls in the first place may also provide a reason *not* to expect much catch up behavior when they are lifted. To the extent that inflation perpetuates itself through self-fulfilling inflationary expectations, *and* to the extent that the shock-treatment of controls succeeds in reducing inflationary expectations, temporary controls could conceivably have a permanent effect on the price level, even with no change in demand management policies. Whether inflationary expectations actually did decline in August 1971, and how

[2]See, for example, Schultze (1975) or Gordon (1973). Other authors have reached the same conclusion.

[3]Raw materials prices also enter marginal costs, and were never controlled (nor could they have been). However, at least in principle, cost increases for these reasons could have been entirely passed through to prices (except during freezes), and thus should not have led to any catch up behavior, except possibly to restore historic margins.

durable this decline proved to be, are hard to know. The well-known Livingston data on inflationary expectations did show a drop of almost 1 percentage point between June 1971 and December 1971, and it seems natural to attribute this to the controls.[4] But the effects of dismantling the controls are impossible to untangle from the Livingston data since decontrol was gradual and since so many other 'shocks' were buffeting inflationary expectations in 1973–1974.

On balance, both the *a priori* theoretical arguments and what we know about how the 1971–1974 controls actually operated seem to point strongly to a post-controls catch up period, though the degree of catch up could be either more or less than the 100 percent catch up depicted in fig. 1. Estimation of the magnitude of the post-controls catch up is one major objective of this paper.

3. The 1971–1974 controls program[5]

The New Economic Policy of August 1971 began with a 90-day freeze on all wages, prices, and rents, with the exception of taxes, raw agricultural products, and mortgage interest. This was followed by Phase II (November 1971 to December 1972), which set broad standards for permissible wage and price increases and relied on self-administered compliance with regulations issued by the Cost of Living Council (COLC). One notable feature of these regulations was that one-for-one percentage pass-through of cost increases to prices was permitted, subject to some profit margin limitations. The Pay Board permitted increases of 5.5 percent in wages, with an extra allowance for certain fringes, and permitted slightly larger increases for workers with relatively small previous increases.

Phase III, which began in January 1973, was intended to be a transitional period providing gradual deregulation. The profit margin constraint for cost-justified price increases of less than 1.5 percent was removed, and larger than standard price increases that were 'necessary for efficient allocation of resources or to maintain adequate levels of supply' became permissible.[6] Events of early 1973 quickly made Phase III look like a failure as increased food prices and cost-justified price increases in other sectors led to a sharp acceleration of inflation. Indeed, critics of the controls program might argue that this was the beginning of the inevitable post-controls catch up inflation. As a result of the acceleration of inflation, a second freeze on prices was imposed from June to August 1973.

Phase IV, following the end of Freeze II, brought with it regulations even more stringent than those of Phase II. For example, cost-justified price

[4]The series cited was created from Livingston's data by Carlson (1977).
[5]For a good detailed discussion of the controls apparatus, see Kosters (1975).
[6]See Kosters (1975, p. 23).

increases were permitted only on a dollar-for-dollar basis rather than a percentage markup basis, which squeezed historic profit margins. Furthermore, the policy of sector-by-sector decontrol was applied in a manner which prevented needed relative price adjustments in several sectors, including steel, petroleum, health, and food.[7] As a result of sector by sector decontrol, only about 12 percent of the *CPI* remained under control when the controls expired on April 30, 1974.

It is apparent from the history of the controls program that any attempt to model controls successfully must account for the changing character of the program over the 1971–1974 period. The controls were not either 'on' or 'off'. Rather, the stringency of regulation in the controlled sector and, especially, the size of the controlled sector varied over time. Furthermore, circumstantial evidence from Phase III suggests the importance of separating catch up inflation in the *uncontrolled* sector from the effects of controls on inflation in the *controlled* sector. Our model reflects these characteristics of the controls program, at least in a crude way.

4. Estimating the impact of controls

To estimate the effects of controls on the price level, an equation or set of equations is required that will generate two price-level paths: one with controls [$\hat{p}(t)$] and the other without [$\bar{p}(t)$]. Two principle techniques have been used to generate the two series needed for this calculation. The first method uses *dummy variables* in a single equation that is estimated over a period including the controls. $\hat{p}(t)$ is generated by the equation with the controls dummy 'on', and $\bar{p}(t)$ is generated by the same equation with the dummy 'off'. The procedure was criticized on a number of grounds by Lipsey and Parkin (1970) and later by Oi (1976). The major difficulties are:

(a) the method is predicated on the assumption that the controls program does not change any of the parameters of the model (except the constant),
(b) any unusual event that happened during the controls period, but is not captured elsewhere in the equation, will be blamed on the controls.

These problems have led most recent investigators of the question to use *post-sample predictions* instead. In this method, the model is estimated only on pre-controls data, and genuine post-sample predictions constitute the series $\bar{p}(t)$. $\hat{p}(t)$ is identified with the actual price level path; so the prediction errors become the estimated effects of controls. While this method avoids problem (a) above, it seems to aggravate problem (b) since it implies

[7]See Kosters (1975, p. 25).

that, were it not for controls, the model would have fit perfectly in every quarter of the controls period.

We therefore advocate, and utilize in this paper, a third method — that of *within-sample predictions.* An empirical wage–price model, which allows controls to alter several parameters, is estimated right through the controls period. Then the model is used to generate *both* the $\hat{p}(t)$ and $\bar{p}(t)$ paths. This technique seems to obviate the difficulties with dummy variables, and allows the equation to have residuals for reasons other than controls. Though it is analogous to the standard procedure by which large-scale econometric models are used to estimate the effects of monetary and fiscal policy, it seems not to have been used heretofore to address the issue of controls.

Robert Gordon's work on controls is probably the best, and also the best known. When Gordon (1975) generated post-sample predictions from a structural price equation explaining the rate of change of a specially-constructed deflator for non-food non-energy final sales, he estimated that controls kept the 1973:3 price level about $3\frac{1}{2}$ percent below what it otherwise would have been. Then prices started to rise faster than they would have in a world without controls, ultimately leaving the price level in 1975:1 1 percent *above* the no-controls path. Two quarters later, the equation was almost precisely on track. However, when he re-estimated the equation with revised data [Gordon (1977)], this picture changed dramatically. Post-sample predictions from the revised equation suggested a downward deflection of the price level of about 2.4 percent by 1972:4, not too different from his 1975 estimate. But then the equation started underpredicting inflation, and by 1975:1 it was underpredicting the price *level* by about 3.7 percent. The difference [$3.7-(-2.4)=6.1$ percent] can hardly be attributed to a post-controls catch up. When Gordon abandoned the post-sample predictions method and used dummy variables instead, his estimate was that controls lowered the price level by 2 percent from 1971:3 to 1972:4, and the post-controls catch up gave this all back between 1974:2 and 1975:1.

McGuire (1976) also used a structural model to make post-sample predictions, but both his specification and his conclusion differed markedly from Gordon's. When he used the Consumer Price Index (*CPI*) to measure inflation, he found that controls lowered the inflation rate *slightly* in 1971:4 through 1972:4, with the maximum effect on the price *level* (about 1.5 percent) coming in 1972:2. Thereafter, controls actually raised the inflation rate, and by 1974:2 the price level was about 3.5 percent *higher* than it would have been without controls. These results are surprising, but some of them must be due to the extraordinary behavior of food and fuel prices, which, in our view, ought not to be blamed on controls. (However, McGuire obtained similar results with the non-farm *GNP* deflator.)

Moreover, McGuire's equation badly underestimates inflation in 1971:1 through 1971:3. Since we look for an effect of controls by looking for *over*predictions starting in 1971:4, this hardly inspires confidence in the results. If we were to add the average prediction error of these three quarters to all the forecasts, we would conclude that controls lowered inflation in the non-farm *GNP* deflator in every quarter from 1971:4 through 1973:3, with a cumulative effect on the price level of 2.5 percent. Thereafter, controls raised inflation so that the 1974:2 price level was only 1.2 percent below what it would have been without controls. This certainly changes the picture drastically, and is consistent with Gordon's estimates — even without making any adjustment for energy.

An earlier study by Lanzillotti, Hamilton and Roberts (1975) used three conventional wage–price models to estimate the effects of Phases I and II. All models used the private non-farm deflator to measure inflation. When the models were estimated through 1971:2 and used to generate post-sample predictions for 1971:3–1972:4, they yielded estimates of the downward deflection of the annual inflation rate averaging 3.6 percent, 1.9 percent, and 0.2 percent. Closer agreement was reached when the models were estimated through 1972:4 and controls appraised by a dummy variable. Estimated downward deflections of the annual inflation rate became 2.0 percent, 1.6 percent, and 1.3 percent. It is interesting that the two methods give such different results in this case.

Other, less conventional, methods have also been used to appraise the Nixon controls — with disparate results.[8] This issue seems unresolved empirically.

5. A new quantitative model of controls

In commenting on Gordon's estimate of the effects of controls, Nordhaus (1975, p. 665) despaired that '. . . the methodology that Gordon and others use to test for incomes policies is inadequate. Can't economists be more creative than to use dummy variables? Why can't we *model* price controls and test the model explicitly?' This section reports on an attempt to do precisely this, though the creativity involved is perhaps not very great.

The model is based on two main ideas. First, that while conventional wage–price models determine the rates of inflation in the controlled and uncontrolled sectors, the parameters of the models may differ between the two sectors. Second, that controls are not strictly a *qualitative* phenomenon, being either 'on' or 'off', but rather have a *quantitative* aspect that has heretofore been ignored. Specifically, the fraction of all prices controlled varied from month to month and can, in principle, be observed.

[8]See, for example, Darby (1976) and Feige and Pearce (1976).

5.1. The price equation

In order to capture more accurately the timing of the shifts from one phase of controls to the next, and also to give more observations during the period of controls, monthly rather than quarterly data were used. However, there is a limited amount of information in the 32 monthly observations on the controlled economy, which in turn limits the number of parameters that can realistically be estimated with these data.

Our basic measure of inflation is the rate of change of the *CPI* exclusive of food and energy prices (π_t^*); we make no attempt here to explain the behavior of food and energy prices. For the period without controls, our model of inflation is quite conventional:

$$\pi_t^* = \alpha_0 + \alpha_1 \Delta_t + \sum_{j=0}^{n} \beta_j w_{t-j} + \sum_{j=0}^{q} \phi_j m_{t-j} + \sum_{j=0}^{m} \gamma_j D_{t-j} + C_t + e_t, \tag{1}$$

where

Δ_t = the deviation of the growth rate of productivity from its trend,
w_t = the rate of change of trend unit labor costs (i.e., money wages deflated by trend productivity levels),
m_t = the rate of change of raw materials prices,
D_t = an indicator of product-market demand pressures, for which two different measures were tried,
C_t = a term for post-controls 'catch up' inflation, to be explained below,
e_t = a white noise disturbance term.

This can be thought of as a standard 'markup' equation.

When controls are in effect, an essentially identical equation is assumed to apply to the *uncontrolled sector* (sector 1):

$$\pi_{1t}^* = \alpha_0 + \alpha_1 \Delta_t + \sum_{j=0}^{n} \beta_j w_{t-j} + \sum_{j=0}^{q} \phi_j m_{t-j} + \sum_{j=0}^{m} \gamma_j D_{t-j} + C_t + e_{1t}, \tag{2}$$

while an analogous equation with different parameters is assumed to apply in the *controlled* sector (sector 2):[9]

$$\pi^*_{2t} = (\alpha_0 + a_0) + (\alpha_1 + a_1)\Delta_t + \sum_{j=0}^{n} (\beta_j + b_j) w_{t-j} + \sum_{j=0}^{q} (\phi_j + f_j) m_{t-j} + \sum_{j=0}^{m} (\gamma_j + c_j) D_{t-j} + e_{2t}. \tag{3}$$

Notice that two sets of identifying assumptions are made: (a) the existence of controls does not alter the parameters of the equation for the uncontrolled sector, and (b) Δ_t, w_t, m_t, and D_t are all the same in the two sectors. While these assumptions can be questioned, it is hard to see what we can do without them. A full wage–price model cannot be estimated from 32 monthly observations.

Since the left-hand side variables in (2) and (3) are unobserved, we need to link the two equations via the following identity:

$$\pi^*_t = \lambda_t \pi^*_{2t} + (1-\lambda_t)\pi^*_{1t} = \pi^*_{1t} + \lambda_t(\pi^*_{2t} - \pi^*_{1t}), \tag{4}$$

where λ_t is the relative importance of the controlled sector. Using (4), (2) and (3) can be combined into a single equation in variables that are actually observed:

$$\pi^*_t = \alpha_0 + a_0\lambda_t + \alpha_1\Delta_t + a_1\lambda_t\Delta_t + \sum_{j=0}^{n} \beta_j w_{t-j} + \lambda_t \sum_{j=0}^{n} b_j w_{t-j} + \sum_{j=0}^{q} \phi_j m_{t-j} + \lambda_t \sum_{j=0}^{q} f_j m_{t-j} + \sum_{j=0}^{m} \gamma_j D_{t-j} + \lambda_t \sum_{j=0}^{m} c_j D_{t-j} + (1-\lambda_t) C_t + e_t, \tag{5}$$

where e_t is white noise if e_{1t} and e_{2t} are. (Possible contemporaneous correlation between e_{1t} and e_{2t} thus becomes irrelevant.)

5.2. *The specification of catch up inflation*

To make (5) operational, we need a specification of the catch up term, C_t, and this is not easy to do in a theoretically satisfying way. What we expect to happen is something like the following. Industry i gets decontrolled in

[9]That there is no catch up term in eq. (3) is definitional since every industry shifts from the controlled sector into the uncontrolled sector when it is released from controls.

month t_i, and at this time has a current price which is g_i percent below its equilibrium price. Call g_i the *disequilibrium gap*. During the next R months we expect super-normal inflation rates for this industry as it 'catches up'. If a set of distributed lag coefficients, v_j^i (where $j=t-t_i$=number of months elapsed since decontrol), shows the time pattern by which industry i returns to its equilibrium price, the additional catch up inflation attributable to industry i would be: $r_i g_i v_j^i$ in month t_i+j $(j=0,1,\ldots,R)$, where r_i is the relative importance of industry i. Now in fact we cannot really identify which industries were decontrolled in which months, nor can we hope to estimate separate disequilibrium gaps and lag patterns by industry. It seems that the best we can do is to invoke the heroic assumptions that (a) all industries catch up with the same distributed lag pattern: $v_0, v_1, \ldots, v_R$; and (b) all industries have the same disequilibrium gap, g, on the date they are decontrolled. The latter is, in fact, an implicit theory of the behavior of the regulators. Under these assumptions, the catch up term can be written

$$C_t = g \sum_{j=0}^{R} v_j \delta_{t-j},$$

where δ_{t-j} is the fraction (relative importance) of the *CPI* that is decontrolled in month t, and g can be identified by the restriction that the v's sum to unity.

5.3. *Controlled relative importance, λ_t*

This time series was constructed from fragmentary information in various publications of the Cost of Living Council (COLC), supplemented by interpolation and some guesswork. Table 1 shows the series, and the following paragraphs explain how it was constructed.

Table 1

Relative importance of controlled prices, λ_t.

Month	1971	1972	1973	1974
January	0	0.818	0.477	0.426
February	0	0.813	0.477	0.426
March	0	0.808	0.477	0.280
April	0	0.802	0.477	0.121
May	0	0.722	0.477	0
June	0	0.698	0.699	0
July	0	0.673	0.912	0
August	0.456	0.649	0.669	0
September	0.912	0.624	0.426	0
October	0.912	0.600	0.426	0
November	0.904	0.576	0.426	0
December	0.845	0.551	0.426	0

According to COLC's first quarterly report,[10] the first freeze covered 91.2 percent of the *CPI*. Since it began on August 15, 1971, we entered half of this amount for August, and the full amount for September and October.[11] The same document reported that 81.8 percent of the *CPI* was controlled on December 31, 1971. Since most of the 9.4 percent drop in coverage came right after the freeze ended on November 15, we guessed at the numbers for November and December as shown in the table.

A controlled relative importance of 80.2 percent on March 31, 1972 was reported in ESPQR (January–March 1972), so this was recorded for April 1972 and February and March were simply interpolated. The small firm exemption occurred in May 1972. While the COLC offered no estimate of the fraction of the *CPI* that was thereby decontrolled, we made a guess of 8 percent based on its report that 13 percent of wages were decontrolled by this action (ESPQR, July–September 1972, p. 4). Then for the period May–December 1972, we simply extrapolated the decontrol rate that had been observed for Phase II between November 1971 and April 1972.

All the remaining observations came from a COLC press release (see ESPQR, January–May 1974) which reported the following values for λ_t:

t	λ_t
January 1, 1973	0.477
June 1, 1973	0.426
September 10, 1973	0.426
March 1, 1974	0.280
April 1, 1974	0.242
April 18, 1974	0.121

The remainder of table 1 was filled out from these numbers and from the assumptions that (a) Freeze II, which included half of June 1973, and half of August 1973, controlled the same fraction of the *CPI* as did Phase I; (b) nothing was decontrolled between January 1, 1973 and June 1, 1973.

From the constructed λ_t series, the δ_t series was created, with one exception, by the following definition:

$$\delta_t = \lambda_{t-1} - \lambda_t \quad \text{if} \quad \lambda_t \leqq \lambda_{t-1},$$

$$\delta_t = 0 \qquad \text{if} \quad \lambda_t > \lambda_{t-1}.$$

[10]U.S. Cost of Living Council, *Economic Stabilization Program Quarterly Report, Covering the Period August 15 through December 31, 1971*. Henceforth these documents are referred to as 'ESPQR' with the appropriate dates.

[11]The BLS notes in its publications that the *CPI* applies to the month as a whole, not to any specific date.

Had there been only one freeze, this definition would have produced a series of positive (or zero) numbers which sum to 0.912. But Freeze II increased λ_t temporarily from 0.477 to 0.912. In principle one would like to specify a second catch up from some new disequilibrium gap caused by the second freeze. This, however, is undoubtedly too much to ask of the limited amount of data, so the actual δ_t equals $\lambda_{t-1}-\lambda_t$ only if $\lambda_{t-1} \leqq \lambda_{t-i}$ for all $i=1,2,\ldots$. This means that the catch up from the first freeze was assumed to continue during the second freeze, and that the second freeze (which lasted only two months) created no new disequilibrium gap of its own.

5.4. Estimation results

The distributed lags in eq. (5) were estimated by the Almon lag technique, after some searching over alternative lag lengths for n, q, m, and R. For wages, materials prices, and demand, a third-degree polynomial with no endpoint constraints was adopted. (There was no experimentation with this choice.) On estimation, materials prices surprisingly turned out to be totally insignificant, and hence were dropped from the equation.

To conserve on parameters, the lag coefficients in the catch up term were at first assumed to follow a quadratic constrained to zero at the far end. We expected to find a lag distribution that rose to a peak and then declined because the COLC often exacted pledges from newly-decontrolled industries that they would exercise restraint in raising prices at first. However, on estimation, this quadratic turned out to be almost exactly linear, so a linear form was substituted.

The productivity deviation variable, Δ_t, was originally entered in distributed lag form. However, preliminary estimates made it apparent that only the contemporaneous value mattered, so the lags were dropped. Similarly, the interaction term, $\lambda_t\Delta_t$, which would indicate any differential effects of productivity deviations under controls, proved insignificant and was dropped from the regression.

Choices of aggregate demand variables are severely circumscribed when monthly data are required. We tried two alternative measures: the logarithm of the unemployment rate, and the residuals from a regression of the logarithm of real personal income on time (henceforth, 'personal income deviations'). More will be said about the choice between these two variables presently.

Finally, eq. (5) as written contains many parameters pertaining only to the controlled sector (all the terms involving the variable λ_t). Not surprisingly, given that λ_t is zero in most quarters of the sample, there was fierce multicollinearity among these variables; so further identifying restrictions were needed before estimation could proceed. These were:

(i) $b_j = 0$ — controls did not alter the pass through of wages into prices;
(ii) $c_j = 0$ — controls did not alter the effect of aggregate demand on prices.

Of these, (i) accords well with the intent of the controls program, but (ii) is a regrettable restriction forced upon us by the weakness of the data. It is well known that the effect of aggregate demand on prices in equations of this type is relatively weak and difficult to estimate. Asking the data to yield two *different* sets of distributed lag estimates was to ask far more than the data could deliver.[12]

With all these simplifications, the equation actually estimated was

$$\pi_t^* = \alpha_0 + a_0 \lambda_t + \alpha_1 \Delta_t + \sum_{j=0}^{n} \beta_j w_{t-j} + \sum_{j=0}^{m} \gamma_j D_{t-j} + g(1-\lambda_t) \sum_{j=0}^{R} v_j \delta_{t-j} + e_t. \tag{6}$$

When eq. (6) was estimated by ordinary least squares using monthly data from January 1961 through December 1977 and the *log unemployment rate* ($\log U_t$) for D_t, the results were as follows:[13]

$$\begin{aligned} \pi_t^* = {} & -\underset{(0.0018)}{0.0065} - \underset{(0.00044)}{0.00223}\lambda_t - \underset{(0.061)}{0.112}\Delta_t + \underset{(0.061)}{0.819}w \\ & - \underset{(0.00057)}{0.00263}\log U + \underset{(0.0123)}{0.0877}(1-\lambda_t)\delta, \end{aligned} \tag{7}$$

$R^2 = 0.768$, D–$W = 2.05$, $S.E.R. = 0.00141$, $n = 204$ months, mean of dependent variable $= 0.0035$.

The reported coefficients of w, $\log U$, and δ are actually the sums of distributed lag coefficients, where the longest lags were 34, 11, and 8 months respectively. Lag lengths were chosen to minimize the sum of squared residuals.

[12] Because the c_j were set to zero, the model implies that the marginal effects of controls did not depend on the state of aggregate demand. Implausibly large estimates of c_j were obtained when this assumption was relaxed. Nonetheless, simulations based on this equation led to estimated effects of controls that were similar to those reported below.

[13] Standard errors are in parentheses. The data were seasonally unadjusted. Eleven monthly dummy variables were included in the regression, but not shown in eq. (7).

The corresponding equation when *personal income deviations* (*PI*) were used to measure demand was

$$\begin{aligned}\pi_t^* = &\underset{(0.0004)}{0.0018} - \underset{(0.00053)}{0.00252}\lambda_t - \underset{(0.055)}{0.238}\Delta_t + \underset{(0.062)}{0.673}w \\ &+ \underset{(0.0060)}{0.0147}PI + \underset{(0.0127)}{0.0623}(1-\lambda_t)\delta,\end{aligned} \tag{8}$$

$$R^2 = 0.763, \quad D\text{–}W = 2.07, \quad S.E.R. = 0.00142.$$

Maximum lag lengths in this equation were 39 for wages, 29 for personal income deviations, and 7 for the catch up.[14]

As the contemporaneous feedback of prices on wages within a single month is quite trivial, and since there was not even a hint of autocorrelation in the residuals, it did not seem worthwhile to use econometric techniques more sophisticated than OLS. The fits of the equations are about the same, and are remarkably good given that the left-hand side variable is a *monthly* inflation rate — an extremely noisy series.

The estimated effects of controls in the two models are similar, though certainly not identical. The coefficients of λ_t imply that a full controls program ($\lambda_t = 1$) — something we never had — would have reduced the annual inflation rate by about 2.6 or 3.0 percentage points, depending on which equation we choose. For a more typical value of λ like 0.6, the implied reductions in the annual inflation rate would be 1.6 or 1.8 percentage points. Continued over a period of 32 months, such a program would lower the price level by about 4.2 or 4.7 percent. However, this crude calculation is not enough to estimate the effects of controls because (a) some catch up begins as soon as the first industry gets decontrolled (which happened in November 1971), and (b) lower inflation rates lead to lower wage settlements, which in turn lead to lower inflation, and so on in a typical wage–price spiral. For both these reasons, a full wage–price model is needed to appraise the effects of controls. Results obtained with such a model are presented in the next section.

The two equations are a little farther apart when it comes to estimating the amount of catch up inflation. Under the assumption that the v_j sum to unity, eq. (7) implies that the typical disequilibrium gap existing when an industry was decontrolled was 8.8%. Eq. (8) estimates the gap to have been only 6.2%. (The estimated length of time it took to rectify this gap was almost the same in the two models.) Thus, on balance, it is clear that eq. (7)

[14]Individual lag coefficients for both equations are available on request.

(with unemployment) gives a more pessimistic view of controls than does eq. (8) (with personal income deviations).[15]

While the aforementioned coefficients are of primary concern in this study, the others are worth a brief mention. The coefficient of the short-run deviation of productivity from trend is negative as expected and significant in both regressions. Not surprisingly, it is smaller in the equation using $\log U_t$ since the unemployment rate presumably captures the short-run labor hoarding from which these short-run productivity deviations partly derive.

The sum of the lag coefficients on trend unit labor costs (and also the shape of the lag distribution) differs between the two equations. In eq. (7), this sum is close to, but significantly different from, the 1.0 value that is often associated with the natural rate. In eq. (8), this sum is only 0.67, which is puzzlingly low.[16]

The estimated effects of aggregate demand on prices, while small, were certainly not negligible. In eq. (7), an increase in the unemployment rate from 5% to 6% would eventually lower the *annual* inflation rate by 0.57 percentage points, which seems in line with other estimates. In eq. (8), where personal income deviations are used to measure demand, a sustained $2\frac{1}{2}\%$ increase of real personal income above trend — which corresponds roughly to a 1 percentage point drop in the unemployment rate — is estimated to add 0.44 percentage points to the annual inflation rate.

5.5. Simulation results

Before we can estimate the effects of controls on the price level, it is necessary to model the feedback of prices into wages. Taking a cue from other investigators who had failed to turn up any direct effect of controls on the wage equation, we estimated the following conventional Phillips curve:

$$\underset{(0.0047)}{W_t = 0.0070} + \underset{(0.174)}{0.965\pi} - \underset{(0.00148)}{0.00082 \log U_t} \tag{9}$$

$$R^2 = 0.64, \quad D\text{–}W = 2.28, \quad S.E. = 0.00388,$$

where W_t is the rate of change of money wages and π is the rate of change of the all-items *CPI*. The coefficients of inflation and unemployment are sums of coefficients of distributed lags of lengths 42 and 4 months, respectively.

[15] While we did not constrain the estimates in any way, there is good consistency between the coefficients of λ and δ. The coefficient of λ implies that the 'disequilibrium gap' for an industry that was controlled for the full 32 months would have been 7.3% or 8.3%, depending on which equation is used.

[16] In this context, it is worth repeating that materials prices were tried, but obtained a coefficient near zero. For those interested in such things, we report that the rate of growth of money (measured by M_2) made no independent contribution to either equation.

The sum of the inflation coefficients is close to unity, and there is remarkably little sensitivity of wage inflation to unemployment.

The distributed lag on past inflation rates in eq. (9) can be given two alternative interpretations. On the one hand, it can measure the catch up process of wages to past (unanticipated) inflation, in which case actual inflation rates are being used as proxies for unanticipated inflation rates. On the other hand, it can be considered a proxy for future anticipated inflation rates. The latter interpretation seems to be more popular, and raises the following question: If, as some proponents of controls claim, the controls program had a direct effect on inflationary expectations, would not the distributed lag proxy be systematically biased during the controls program? Were this true, it seems to us, a controls variable should 'sop up' this systematic misestimation of expected inflation when added to eq. (9). This proved not to be the case.

A word on identification should also be entered. The price equation is identified from the wage equation by the omission of lagged prices from the former and of lagged wages, the productivity term, and the controls variables from the latter. (In addition, one version of the price equation uses a different demand variable than the wage equation.) These identifying restrictions were tested. Neither productivity deviations nor λ_t entered (9) significantly.[17] Lagged prices were insignificant when added to eq. (8), but did contribute marginally to eq. (7).[18] Lagged wages were significant ($F=4.1$) when added to eq. (9), but produced implausible parameter estimates.[19]

To close the model, it was necessary to append the following two identities:

$$w_t \equiv W_t - \rho_t, \tag{10}$$

$$\pi_t \equiv r_t^* \pi_t^* + r_t^F \pi_t^F + r_t^E \pi_t^E, \tag{11}$$

where

ρ_t = the trend rate of increase of productivity applicable to month t (see data appendix for a description),
r_t^* = the relative importance of non-food non-energy items in the *CPI*,
π_t^F = the rate of inflation of the *CPI* food component,
π_t^E = the rate of inflation of the *CPI* energy component,
r_t^F = the relative importance of food items in the *CPI*,
r_t^E = the relative importance of energy items in the *CPI*.

[17]A 'catch up' variable similar to that used in the price equation was significant. However, since no direct depressing effect of controls on W_t was found, we judged this correlation to be spurious.

[18]The F ratio was 3.2, which is significant at the 5% level, but not at the 1% level.

[19]The sum of the price coefficients was 4.1 while the sum of the wage coefficients was −2.8.

Tables 2 and 3 display the results obtained in a 77-month dynamic simulation of two versions of this four-equation model, one using (7) and the other using (8). The simulation used actual wage and price data up to July 1971, and predicted values thereafter, with no error correction. Not shown in either table, but worthy of note, is the fact that both models track history extremely well when controls are 'on', and are almost precisely on track when the simulation ends in December 1977. The tables also include, in parentheses beneath each estimate, approximate standard errors derived from a Monte Carlo procedure.[20]

Table 2

Effects of controls on the non-food, non-energy Consumer Price Index, 1971–1975, based on Model 1 (unemployment rate), in percent.

Month	1971	1972	1973	1974	1975
January	—	−1.031	−1.986	−2.986	+0.183
		(0.225)	(0.618)	(0.988)	(1.312)
February	—	−1.187	−1.981	−3.078	+0.171
		(0.263)	(0.647)	(1.015)	(1.322)
March	—	−1.344	−2.002	−2.963	+0.161
		(0.300)	(0.676)	(1.040)	(1.331)
April	—	−1.501	−2.045	−2.565	+0.153
		(0.338)	(0.705)	(1.065)	(1.341)
May	—	−1.600	−2.108	−1.939	+0.147
		(0.371)	(0.734)	(1.098)	(1.350)
June	—	−1.688	−2.247	−1.400	+0.143
		(0.404)	(0.770)	(1.134)	(1.360)
July	—	−1.765	−2.462	−0.938	+0.141
		(0.436)	(0.817)	(1.170)	(1.371)
August	−0.101	−1.834	−2.618	−0.556	+0.141
	(0.021)	(0.467)	(0.853)	(1.206)	(1.381)
September	−0.304	−1.895	−2.674	−0.253	+0.143
	(0.064)	(0.499)	(0.880)	(1.238)	(1.392)
October	−0.506	−1.947	−2.744	−0.031	+0.146
	(0.107)	(0.529)	(0.907)	(1.265)	(1.403)
November	−0.706	−1.990	−2.819	+0.112	+0.151
	(0.149)	(0.560)	(0.934)	(1.286)	(1.414)
December	−0.876	−2.024	−2.900	+0.174	+0.157
	(0.188)	(0.590)	(0.961)	(1.301)	(1.425)

[20]The estimated effects of controls in each month are very complicated non-linear functions of all the parameters of the model. Their sampling distributions are unknown. We estimated standard errors as follows. First, we made 100 drawings from the (known) joint normal distribution of the regression coefficients. For each drawing, a 77-quarter simulation of the type described in the text was run. From these 100 runs, a mean and a standard deviation of the estimated effects of controls was calculated for each month. The standard deviations appear in tables 2 and 3.

Table 3

Effects of controls on the non-food, non-energy Consumer Price Index, 1971–1975, based on Model 2 (personal income deviations), in percent.

Month	1971	1972	1973	1974	1975
January	—	−1.180 (0.273)	−2.616 (0.713)	−4.066 (1.133)	−2.242 (1.428)
February	—	−1.365 (0.318)	−2.662 (0.742)	−4.193 (1.166)	−2.269 (1.439)
March	—	−1.550 (0.362)	−2.729 (0.772)	−4.150 (1.192)	−2.294 (1.450)
April	—	−1.737 (0.407)	−2.814 (0.802)	−3.880 (1.214)	−2.317 (1.461)
May	—	−1.872 (0.445)	−2.916 (0.834)	−3.434 (1.239)	−2.338 (1.472)
June	—	−1.999 (0.482)	−3.089 (0.875)	−3.056 (1.270)	−2.355 (1.483)
July	—	−2.117 (0.518)	−3.336 (0.930)	−2.746 (1.303)	−2.371 (1.494)
August	−0.115 (0.026)	−2.229 (0.553)	−3.524 (0.973)	−2.506 (1.335)	−2.384 (1.505)
September	−0.343 (0.078)	−2.331 (0.588)	−3.619 (1.003)	−2.335 (1.364)	−2.395 (1.516)
October	−0.571 (0.130)	−2.425 (0.621)	−3.721 (1.035)	−2.233 (1.387)	−2.404 (1.527)
November	−0.797 (0.181)	−2.512 (0.653)	−3.830 (1.067)	−2.200 (1.404)	−2.412 (1.538)
December	−0.995 (0.228)	−2.593 (0.685)	−3.945 (1.099)	−2.212 (1.417)	−2.417 (1.548)

The two models tell somewhat different stories about the effects of controls. As expected, Model 2, with personal income deviations as the demand variable, is more optimistic about controls than is Model 1, which uses the unemployment rate.

According to Model 1, controls reduced the non-food non-energy inflation rate by about $1\frac{3}{4}$ percentage points during their first year, and then had only a small effect on inflation until Freeze II (June–August 1973) began. This freeze had a noticeable effect, however, and by February 1974, when the maximum effect of controls apparently was felt, non-food non-energy prices were about 3.1 percent lower than they would have been without controls.

At this point, extra catch up inflation in the uncontrolled sector began overwhelming whatever effects controls were still having in the controlled sector, and by October 1974 the price level was just about where it would have been had controls never been put in place. According to this model, catch up was particularly severe between March and September of 1974, and there was virtually no permanent effect on the price level.

Model 2 agrees with this assessment in broad outlines, but differs substantially in its details. According to this model (which measures demand by personal income deviations), controls chopped about 2 points off the inflation rate between August 1971 and July 1972, and pushed prices down almost another percent between July 1972 and June 1973. Freeze II again shows up clearly, and the ultimate downward deflection of the price level — which, once again, comes in February 1974 — is 4.2%. Up to this point, while Model 2 is more optimistic about controls, the predictions of the two models are within one standard deviation of each other.

More disagreement emerges in the catch up period. While both agree that catch up started in March 1974, and was strongest between March and August, Model 2 estimates the catch up to have been far from complete. According to this model, the price level in November 1974 remained 2.2 percent below what it would have been in the absence of controls, and was *permanently* lowered by about 2.4 percent. These last two estimates differ from those of Model 1 by about 1.7 standard deviations.

Our conclusions about the catch up would therefore be firmer if one model clearly dominated the other. However, such is not the case. In terms of within-sample fit, Model 1 does slightly better. A variety of genuine post-sample prediction tests that we ran resulted in a draw: each model outperformed the other in some periods. In terms of 'reasonableness' of the coefficients, Model 1 seems to have a more sensible wage elasticity of the price level than does Model 2. On balance, Model 1 perhaps has a small edge, but the margin is slight.

6. Conclusions

Armed with the estimates from section 5, let us now address the questions asked in the opening section: by how much did controls hold down prices while they were in effect? how much catch up inflation was there when controls were lifted? to what extent can decontrol account for the acceleration and deceleration of inflation in 1973–1975?

As we have already seen, the two models differ modestly in their answer to the first question. In both cases, the maximal estimated effect of controls on the price level came in February 1974. But in Model 1 (based on the unemployment rate) this was estimated to have been −3.1 percent while in Model 2 (based on personal income deviations) the estimate was −4.2 percent.

Both models agree that there was catch up inflation after February 1974, though they differ on the amount. Focusing on the period from February to October 1974, Model 1 estimates that catch up carried the price level up 3 percent while Model 2 makes this only 2 percent.

These disagreements, while not trivial, are perhaps not that great. Reasonably good agreement is also reached on the third question, as table 4 illustrates. The first two lines summarize the behavior of the *CPI* stripped of food and energy prices — the variable called π^* in the model. Examination of this series shows a clearly defined period of 'double digit' inflation lasting from February to October 1974, with much lower inflation rates both before and after. Because this period of peak inflation was 8 months long, 'pre-peak' and 'post-peak' periods of 8 months' duration were chosen for symmetry. As can be seen in line 2, the non-food non-energy inflation rate accelerated almost 7 percentage points in early 1974 and decelerated almost 5 percentage points in late 1974. How much of this roller-coaster pattern of inflation can be traced to the controls?

Table 4

Acceleration and deceleration of inflation, 1973–1975.

	Pre-peak period (June 1973–Feb. 1974)	Peak period (Feb.–Oct. 1974)	Post-peak period (Oct. 1974–June 1975)
Actual data			
(1) Annual rate of inflation (seasonally adjusted)	5.90%	12.72%	7.84%
(2) Acceleration (+) or deceleration (−) from previous period	—	+6.82%	−4.88%
Contribution of wage–price controls according to Model 1			
(3) To inflation rate	−1.34%	+5.12%	+0.28%
(4) To acceleration/deceleration	—	+6.46%	−4.84%
Contribution of wage–price controls according to Model 2			
(5) To inflation rate	−1.82%	+2.17%	−0.20%
(6) To acceleration/deceleration	—	+3.99%	−2.37%

Lines (3)–(6) give the answers according to the two models. Lines (3) and (5) come directly from the estimated effects of controls on the *price level* in tables 2 and 3 by converting the estimated effects on price *levels* into effects on the *annual inflation rate*. These lines merely repeat the differences between the two models that we have already noted. Lines (4) and (6) indicate how much of the acceleration and deceleration of inflation each model 'blames' on the controls. The conclusion from Model 1 is that almost all of the movement in non-food non-energy inflation during this two-year period was a result of the controls. Model 2, with its incomplete catch up, blames only about half of the acceleration and deceleration on controls.

A more detailed look at the period from June 1971 to June 1975 as a whole reveals that, while the models disagree moderately on the effects of controls on the *mean* inflation rate, they agree very well that controls added substantially to the *variance* of the inflation rate. Table 5 displays this conclusion. It is constructed by dividing the period June 1971–June 1975 into eight half-years and computing for each the annual rate of inflation (a) that actually occurred in non-food non-energy prices, (b) that would have occurred in the absence of controls according to Model 1, (c) that would have occurred in the absence of controls according to Model 2. The differences in the means simply repeat what we have already seen. But the differences in the standard deviations are strikingly large in both models.

Table 5

Effects of controls on the mean and standard deviation of inflation,[a] June 1971–June 1975.

	Mean inflation rate[b]	Standard deviation of inflation rate
(a) Actual data	5.91 %	3.70 %
With no controls according to:		
(b) Model 1	5.63 %	1.44 %
(c) Model 2	6.38 %	1.95 %

[a]For non-food non-energy *CPI*.
[b]Based on annual rates of inflation over 8 half-year periods.

The models agree that controls at least doubled the standard deviation of the inflation rate over the four-year period as a whole. If, as we suspect may be the case, the most severe social costs from inflation are attached to its variance, rather than to its mean, this may be the most serious condemnation of controls.

Data appendix

Prices

The rates of change of the *CPI* (π_t) and its non-food non-energy components (π_t^*) were computed from official BLS data on the *seasonally unadjusted levels* of these series. Seasonal adjustment was accomplished by adding 11 monthly dummies to the estimated equations.

Materials prices

Materials prices were measured by the Producer Price Index for crude materials for further processing.

Wages

The rate of change of money wages was measured using the BLS's series on average hourly earnings of production or non-supervisory workers on manufacturing payrolls, seasonally unadjusted and corrected for overtime and interindustry shifts. Monthly changes were used with no attempt to 'smooth' the data.

Productivity

Quarterly BLS data on the growth rate of output per hour in non-farm private business over the period 1947 to 1978 were used to estimate the following regression:

$$\underset{}{G_t} = \underset{(0.00100)}{0.00668} - \underset{(0.00189)}{0.00167} D1_t - \underset{(0.00242)}{0.00335} D2_t,$$

$$R^2 = 0.046, \quad D\text{–}W = 1.96, \quad S.E. = 0.00866,$$

where G_t is the quarterly growth rate of productivity, $D1_t$ is a dummy which is one beginning in 1966:1, and $D2_t$ is a dummy which is one beginning in 1973:2.[21] The fitted values from this regression were used to generate the 'trend productivity' factor (called ρ_t in the model) by phasing in the two discrete drops in the productivity trend linearly over 12-month periods centered on 1966:1 and 1973:2. The differences $G_t - \rho_t$ constituted the variable Δ_t used in the regressions. Both ρ_t and Δ_t as just defined are *quarterly* series. They were converted to *monthly* series by linear interpolation.

Personal income deviations

The personal income deviations series was created by deflating personal income by the Consumer Price Index, and then fitting the following regression to monthly data from January 1957 through December 1978:

$$\log(\text{real personal income}) = \underset{(0.003)}{0.201} + \underset{(0.00002)}{0.00331}\ \text{time},$$

$$R^2 = 0.988, \quad D\text{–}W = 0.03.$$

[21]These break points were chosen to minimize the sum of squared residuals in the regression. It is interesting that they are nonetheless statistically insignificant.

Unemployment

The unemployment rate was measured by the BLS's unemployment rate for all civilian workers.

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