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MONEY AND PRICE DISPERSION IN THE UNITED STATES

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ABSTRACT

This paper reports an empirical test of a price dispersion equation, using data on the U.S. after World War II. The equation, derived elsewhere from a version of the partial information-localized market models, relates price dispersion to the magnitude of changes in the aggregate disturbances. In order to test the model a series on price dispersion is computed using annual wholesale price indexes for the period 1948-76. The data on money shocks are the unanticipated money growth series estimated by Barro. The tests also include a measure of aggregate-real disturbances.

From the theoretical point of view the results are negative. They reject the hypothesis that unexpected money shocks, as measured by Barro, affect price dispersion in the way predicted by the model. In a previous paper, a similar model was tested with data from the German hyperinflation and found supported to a considerable extent. The difference in the results may be related to the extreme magnitude of the monetary disturbances during that period, and to the apparently important effect of unincluded relative disturbances in the United States.

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This paper reports an empirical test of a price dispersion equation, using data on the U.S. after World War II. The equation is derived in Hercowitz (1979) from a theoretical framework that is a version of Lucas's (1973) and Barro's (1976) partial information-localized market models.

In this extension, which is briefly summarized in section I, each commodity is characterized by a particular excess demand elasticity. The different reactions to money shocks across markets imply a testable dispersion equation in which the explanatory variables are the magnitude of changes in the aggregate disturbances. The data on money shocks used in the tests are the unanticipated money growth series estimated by Barro (1978). The series on price dispersion is computed using annual wholesale price indexes for the period 1948/76. Vining and Elwertowski (1976) have performed similar computations with a more extensive disaggregation, but using the unweighted price changes. The procedure adopted here in the calculations is reported in section II.

Using these figures, the dispersion equation is tested in section III. From the theoretical point of view the results are negative. They reject the hypothesis that unexpected money shocks , as measured by Barro (1978), affect price dispersion via the channel described in section I. In my previous paper, a similar model was tested with data from the German hyperinflation and found supported to a considerable extent. The difference in the results may be related to the extreme magnitude of the monetary disturbances during that period, and to the apparently important effect of real factors--which directly affect relative excess demands--in the United States.

The results in this paper are also not in line with previous findings related to price dispersion in the U.S. Vining and Elwertowski (1976) show that the variance of relative prices tends to increase with the magnitude of changes in the rate of inflation.[1] Parks (1978) tested a model in which price change dispersion depends on the magnitude of changes in real income and unexpected inflation, which is measured by the changes in the inflation rate (minus a constant).[2] Parks' empirical results strongly support his model.

In the theoretical framework used here, unanticipated inflation is generated by unexpected money or aggregate-real shocks. These disturbances also effect relative prices and thus this model offers a rationalization of the previous findings in terms of the effects of exogenous shocks. Since, empirically, the link between money disturbances and price dispersion is found absent, the explanation offered by the model is not supported by the data. The additional variable considered, the magnitude of changes in a measure of aggregate-real shocks, seems to have a positive correlation with price dispersion.

In any event, the apparent regularity of the phenomenon that price dispersion is related to the magnitude of changes in the inflation rate still lacks a satisfactory explanation in terms of the effects of exogenous shocks or policy variables.

I. Review of the model.

This section contains a short outline of the model described in section I of my previous paper. The only difference introduced here is an aggregate supply shock that affects equally all markets.

The central feature in the localized market models of the type used here, is the combination of competitive markets--in which demand and supply are functions of the perceived relative price--and partial current information. Agents lack knowledge about prices prevailing contemporaneously in other markets, and observe only the local price. The information set available to the agents contains also lagged variables and some incomplete knowledge about factors governing the current period money growth.

In this model each location is interpreted as the market of a specific commodity, and it is characterized by a particular relative price elasticity of excess demand. With different elasticities across markets, unperceived aggregate disturbances have a different impact on each commodity price, and thus generate dispersion of prices. The supply and demand for a particular commodity are aggregations of individual supplies and demands by a large number of agents located in period t in market z. These functions are given in log-linear terms by

$$(1)\quad y_t^s(z) = \alpha^s(z)[P_t(z) - EP_t] + v_t + \varepsilon_t^s(z)$$

$$(2)\quad y_t^d(z) = -\alpha^d[P_t(z) - EP_t] + [M_t - EP_t] + \varepsilon_t^d(z)$$

$P_t(z)$ is the (log of the) nominal price of commodity z and EP_t is the expectation of the average price level, formed conditional on the information available to the local agents. In the supply function, $\alpha^s(z)$ is the relative price elasticity of supply, which is seen as depending on the particular production process of commodity z. v_t represents real aggregate supply shocks. This term can be thought of as, for example, changes in technology or in the availability of resources such as energy. v_t is assumed to be normally distributed with zero mean and variance σ_v^2 . One could specify, as in Barro (1976), that real shocks follow a random walk; then, the innovation in the process would have the same effect as v_t here.

In the demand function M_t is the (log of the) money stock at time t, and the term $M_t - EP_t$ represents a real balance effect. α^d is the relative price elasticity of demand, which is constant across markets. The terms $\varepsilon_t^s(z)$ and $\varepsilon_t^d(z)$ represent relative shifts of supply and demand. The excess demand shift $\varepsilon_t(z) \equiv \varepsilon_t^d(z) - \varepsilon_t^s(z)$ is assumed normally distributed with zero mean and variance σ_ε^2.

At the beginning of the period, new money is introduced into the economy, equally dispersed across markets. Within each market, however, it is randomly distributed among the large number of agents located there. The money stock grows according to $m_t = M_t - M_{t-1} = g_t + \tilde{m}_t$. g_t is known economy wide, and $\tilde{m}_t$ is a currently unobservable random term of zero mean and variance σ_m^2.

From (1) and (2), market clearing in z implies

$$(3)\ P_t(z) = [1 - \lambda(z)]\ EP_t + \lambda(z)[M_t - v_t + \varepsilon_t(z)]$$

where $\lambda(z) \equiv 1/[\alpha^s(z) + \alpha^d]$ --i.e., it is the inverse of the price elasticity of excess demand for z. Over time, each market has a constant $\lambda(z)$; but across markets, $\lambda(z)$ has the average value of λ and the "variance" σ_λ^2.

Following Lucas (1973) and Barro (1976), the solution of the model is obtained using the method of undertermined coefficients. The final expression for $P_t(z)$ in terms of the exogenous variables is

$$(4)\ P_t(z) = M_{t-1} + g_t + \frac{\sigma_m^2 + \sigma_v^2 + \lambda(z)(1/\lambda)\sigma_\varepsilon^2}{\sigma_m^2 + \sigma_v^2 + (1/\lambda)\sigma_\varepsilon^2}[\tilde{m}_t - v_t + \varepsilon_t(z)]$$

The actual relative price of commodity z is calculated by subtracting from $P_t(z)$ the average price across markets

$$(5)\ P_t(z) - P_t = (1 - \theta)\tilde{\lambda}(z)\ (\tilde{m}_t - v_t) + [\theta + \lambda(z)\ (1 - \theta)]\varepsilon_t(z)$$

where $\tilde{\lambda}(z) = \lambda(z) - \lambda$ and $\theta = \dfrac{\sigma_m^2 + \sigma_v^2}{\sigma_m^2 + \sigma_v^2 + (1/\lambda)\sigma_\varepsilon^2}$

The aggregate excess demand shock $\tilde{m}_t - v_t$ appears in (5) multiplied by $\tilde{\lambda}(z)(1 - \theta)$. The fraction $(1 - \theta)$ indicates the part of this disturbance that is perceived locally as a relative shift. Since the supply reaction depends on $\alpha^s(z)$, the smaller this elasticity,--the higher $\tilde{\lambda}(z)$--the larger the price response. This is the channel by which aggregate shocks disperse individual prices. On the other hand, monetary movements that are fully perceived, g_t, are not confused for relative shifts, and therefore they inflate equally all prices.

The variance of relative prices can be calculated by averaging (5) across markets. However, it is empirically more useful to use the variance of the rate of change in individual prices. Prices may disperse over time according to different long-run trends. Since this model focuses on short run distortions caused by imperfect information, one can minimize the effects of these secular price movements by looking at the variance of the rate of change. This variance is defined here as $\gamma_t^2 \equiv (1/N)\ \Sigma\ [P_t(z) - P_{t-1}(z) - (P_t - P_{t-1})]^2$ where N is the total number of markets. γ_t^2 is computed using the difference between (5) and the equivalent for t-1. The resulting expression is

$$(6)\quad \gamma_t^2 = \{(1-\theta)^2\sigma_\lambda^2 + 2[\theta + \lambda(1-\theta)]^2\}\ \sigma_\varepsilon^2 + (1-\theta)^2\sigma_\lambda^2\ [(\tilde{m}_t - \tilde{m}_{t-1}) - (v_t - v_{t-1})]^2$$

This equation is tested in section III. Given σ_m^2, σ_v^2, σ_ε^2 and the disparity in elasticities, represented by σ_λ^2 , equation (6) predicts a positive effect of the magnitude of changes in money and real-aggregate shocks on price change dispersion.

Section II. Computation of the price dispersion measure: U.S. 1948-1976

A price dispersion series is constructed using annual average wholesale price data contained in the Bureau of Labor Statistics WPI subfile tape. The computation has three steps: 1) calculation of the rates of changes in individual prices between each pair of consecutive years in the sample, 2) calculation of the weight that each commodity has in the total value of sales in primary markets, and 3) computation of the average and variance of the price changes using the individual rates of change and weights calculated in 1) and 2). The computed values are shown in columns (1) and (2) of table I.

Individual commodities are defined here as the subgroups of items comprising the WPI. There are 86 of these subgroups at the beginning of the sample in 1948, and 97 of them at its end in 1976. This disaggregation is obtained using the three digit code indexes in the WPI subfile tape. A report of the procedure adopted in the computations is given in more detail in the appendix.

Whereas the problem of the WPI data--discussed by Stigler and Kindahl (1970)--that they do not always reflect discounts from list prices and other terms of trade, it seems that the use of rates of changes in annual average data diminishes considerably this possible error of measurement. The discrepancy between the WPI indexes and those constructed by Stigler and Kindahl is, in general, materially less pronounced when the time unit is a year than when it is shorter. This discrepancy is particularly small in the rates of price change of metals and fuels, which, in terms of their relative importance, are the bulk of the commodities that they consider.

Section III. Empirical Results

This section reports an attempt to test equation (6) using U.S. data for the period 1948-1976. The $\tilde{m}_t$ and v_t series employed in the tests are the residuals in money growth and output equations respectively, as estimated by Barro (1978). In that study, money growth was regressed on a number of current and lagged variables that were considered relevant to predict it. Accordingly, the residuals in the equation were interpreted as the unexpected money growth shocks. The inclusion of a contemporaneous variable--related to government spending[3]--suggests that the residuals could also be interpreted as currently unperceived money, which is the monetary variable that is theoretically relevant here.[4]

Barro then regressed output on current and past values of the estimated money shocks. The residuals in this equation are considered here as the nonmonetary or real disturbances, and are used as values of the v_t variable.[5] In general, however, these residuals are not equal but proportional to v_t. In the framework of this model, the coefficient of v_t in the aggregate output equation is computed by solving for output in z and then averaging across markets. The resulting expression for the (log of the) geometric average output in the economy is:

$$y_t = (1 - \theta)\ (1 - \alpha^d\lambda)\tilde{m}_t + [\theta + \alpha^d\lambda(1 - \theta)]v_t$$

The coefficient of v_t is positive and less than one since $\alpha^d\lambda < 1$.[6] It can be observed that this inequality also implies that monetary shocks have a positive effect on output. According to this equation, the residuals in the regression of output on money shocks are interpreted as the real aggregate shocks times the coefficient $\theta + \alpha^d\lambda(1 - \theta)$. Note that this interpretation relies on an implicit assumption made in the specification of the model that real aggregate shocks affecting the demand side are of minor importance relative to the supply shocks, and therefore can be neglected.

Now we can proceed to test equation (6). For convenience it is rewritten here in the following form

$$(7)\quad y_t^2 = a_0 + a_1(\tilde{m}_t - \tilde{m}_{t-1})^2 - a_2(\tilde{m}_t - \tilde{m}_{t-1})\ (v_t - v_{t-1}) + a_3(v_t - v_{t-1})^2$$

where

$$(8)\qquad a_0 = \{(1 - \theta)^2\sigma_\lambda^2 + 2[\theta + \lambda(1 - \theta)]^2\}\sigma_\varepsilon^2$$

$$a_1 = \tfrac{1}{2}a_2 = a_3 = (1 - \theta)^2\sigma_\lambda^2$$

The variances included in the coefficients are considered constant during the sample period. Shifts in these variances could be another source of variation in γ_t^2. Some reference to this possibility is made below.

As expressed in equation (7), the model predicts that the absolute value of the changes in both monetary and real shocks will positively affect dispersion. This effect is attributable to unequal quantity and price response to unexpected shocks across markets. From the point of view of this hypothesis, the empirical results are in general unsatisfactory and especially negative with respect to the monetary disturbances. The coefficient of the $(\tilde{m}_t - \tilde{m}_{t-1})^2$ variable has the 'wrong' sign in all the equations estimated, although the estimated coefficients are also insignificantly different from zero. On the other hand, the real shock variable appears significantly to affect dispersion in the way predicted by the model.

The estimated equation using the γ_t^2 series computed above is

$$\text{(9)} \quad \gamma_t^2 = \underset{(.001)}{.003} - \underset{(.020)}{.012}\,(\tilde{m}_t - \tilde{m}_{t-1})^2 + \underset{(.026)}{.052}\,(\tilde{m}_t - \tilde{m}_{t-1})\,(v_t - v_{t-1})$$

$$+ \underset{(.019)}{.050}\,(v_t - v_{t-1})^2$$

$R^2 = .32$ D.W. = 1.2 29 observations

where the numbers in parenthesis are the standard errors of the coefficients.

Adding the lagged variables $(\tilde{m}_{t-1} - \tilde{m}_{t-2})^2$, $(\tilde{m}_{t-1} - \tilde{m}_{t-2})\,(v_{t-1} - v_{t-2})$ and $(v_{t-1} - v_{t-2})^2$, yields insignificant coefficients without materially changing the estimated coefficients and standard errors of the contemporaneous variables. Given the low D.W.-statistic, the equation was reestimated using the Cochrane-Orcutt technique without changing the main results in equation (9).[7]

According to the theory, the coefficients of $(\tilde{m}_t - \tilde{m}_{t-1})^2$ and $(v_t - v_{t-1})^2$ must be the same. However, as mentioned before, there is a presumption that the values of the v_t variable contain a less-than-one factor of proportionality, and thus, the coefficient of v_t is likely to be biased upwards.

Additional tests were performed using three alternative measures of price change dispersion. The first is a series constructed by Parks (1978). This dispersion measure was computed using a 12-component breakdown of Personal Consumption Expenditure deflators and weights data, and it is denoted here by $(\gamma_t^2)_{PCE}$. The sample in this regression is 1948-1975.

$$(\gamma_t^2)_{PCE} = \underset{(.002)}{.002} - \underset{(.245)}{.059}\,(\tilde{m}_t - \tilde{m}_{t-1})^2 + \underset{(.322)}{.794}\,(\tilde{m}_t - \tilde{m}_{t-1})\,(v_t - v_{t-1})$$

$$+ \underset{(.265)}{1.000}\,(v_t - v_{t-1})^2$$

$R^2 = .44$ D.W. = 1.9 28 observations

The results here are essentially similar to those obtained in the first regression.

The additional two measures of price change dispersion used are the ones computed by Vining and Elwertowski (1976). From the individual items indexes comprising the Wholesale Price Index, they constructed series of the sample standard deviations of the unweighted individual price changes for the period 1948-1974. This series is denoted here by $(\gamma_t^2)_{WPI}$. The number of items entering each year in the computation ranges from 1159 to 2033. They performed also similar computations using the Consumer Price Index components. The number of items considered for each year ranges from 110 to 311. The symbol used here for this series is $(\gamma_t)^2_{CPI}$.

The regression results using these two measures of γ_t^2 are as follows

$$(\gamma_t^2)_{WPI} = \underset{(.002)}{.009} - \underset{(.027)}{.011} (\tilde{m}_t - \tilde{m}_{t-1})^2 + \underset{(.036)}{.065} (\tilde{m}_t - \tilde{m}_{t-1}) (v_t - v_{t-1})$$

$$+ \underset{(.029)}{.054} (v_t - v_{t-1})^2$$

$R^2 = .22$ D.W. = 1.2 27 observations

$$(\gamma_t^2)_{CPI} = \underset{(.001)}{.003} - \underset{(.009)}{.002} (\tilde{m}_t - \tilde{m}_{t-1}) + \underset{(.012)}{.017} (\tilde{m}_t - \tilde{m}_{t-1}) (v_t - v_{t-1})$$

$$+ \underset{(.010)}{.021} (v_t - v_{t-1})^2$$

$R^2 = .12$ D.W. = .9 27 observations

All the equations estimated yield approximately the same results although with differences in the goodness of fit. These results amount to a rejection of the hypothesis that the magnitude of changes in monetary shocks, as measured by Barro (1978), positively affects dispersion of relative prices through the particular channel described in section I. From the standpoint of this hypothesis, a particularly negative piece of evidence is the monetary contraction in 1960. In this year, the value of $(\tilde{m}_t - \tilde{m}_{t-1})^2$ is the highest in the sample, at the time that price dispersion is extremely low--as it continues to be during the early sixties.

With respect to the real shocks , the evidence seem to some extent to support the hypothesis implied by the model. However, even with this result some caution is in order. If one believes that real shocks hit different markets unequally, they could cause dispersion also or perhaps only by shifting the current σ_ε^2. For example, the oil price shock in late 1973 and 1974 affects some sectors of the economy more than others. Namely, the heavier the use of energy in the production of a particular commodity, the greater the effect of the oil price increase on that industry. Therefore, besides the aggregate impact of this disturbance, it also represents an intensification of the relative differences of excess demand among the markets, with the result of a larger price change dispersion.[8]

The importance of real-relative disturbances that directly affect the magnitudes of relative excess demands is suggested by the observations corresponding to 1974 and the years that follow, by 1973--a year of a sharp foreign demand increase for food products--and by the Korean War. These types of factors are considered neither in the theoretical model nor directly in the empirical analysis. However, if the constructed monetary shock series is independent of these factors, the present procedure for estimating monetary effects on dispersion would still be satisfactory.

IV. Summary and Conclusions:

The aim of this paper was to test with U.S. data a model of price dispersion in which monetary and real-aggregate shocks are the explanatory variables. However, the effect of money shocks that the model predicts--on which this paper is mainly concerned--is not found in the data. Possibly, this effect is of small magnitude relative to that of apparently important excluded variables, and thus is difficult to detect in the present empirical framework. This conjecture is based on a comparison of these results with those from the German hyperinflation, as reported in Hercowitz (1979). In that period, in which money growth is a predominant factor, a dispersion model of the type used here receives substantial support.

In this version of Lucas's (1973) and Barro's (1976) models, money shocks affect both relative prices and aggregate output. The effect on output, well known from those models, occurs because of incomplete information. The additional effect on price dispersion, results from the interaction between incomplete information and different supply elasticities across markets. Empirically, Barro (1977, 1978) found that unanticiapted money has a strong effect on unemployment and output. Using the same monetary variable, the additional hypothesis derived from this extension is rejected. This result is reconciled with Barro's if all markets have the same price elasticity. In this case, the effect on output remains, but that on price dispersion is absent. However, the assumption of equal short run price elasticities in all sectors seems fairly strong.

The empirical result that unanticipated money is not important for price dispersion contrasts with findings of Mills (1927), Vining and Elwertowski (1976), and Parks (1978), which can be summarized as isolating a positive correlation between price dispersion and the magnitude of changes in the inflation rate. A variable of this type is interpreted in Parks' model as the magnitude of unexpected inflation.[9] In the model used here, a correlation between price dispersion and unexpected inflation follows from the effects of money (and real-aggregate) shocks on both. However, since the hypothesis about the effect of money shocks on dispersion is rejected, a link between the theoretical model used here and the previous studies fails to be established.

Table I

Average Price Change, ΔP_t, and Variance of Relative Price Changes, γ_t^2

Wholesale Prices, United States 1948-1976.

Year	ΔP_t (1)	γ_t^2 (2)	Number of commodities* (3)
1948	.080	.0043	64
1949	-.050	.0087	63
1950	.036	.0024	64
1951	.110	.0022	64
1952	-.020	.0053	62
1953	-.013	.0043	64
1954	.002	.0025	62
1955	.005	.0043	61
1956	.039	.0019	63
1957	.033	.0029	61
1958	.017	.0030	64
1959	.003	.0020	66
1960	.000	.0008	66
1961	-.004	.0008	69
1962	.004	.0005	69
1963	-.003	.0012	70
1964	.002	.0010	69
1965	.020	.0018	69
1966	.035	.0013	72
1967	.004	.0017	74
1968	.024	.0012	75
1969	.039	.0015	75
1970	.037	.0018	76
1971	.034	.0019	76
1972	.045	.0025	74
1973	.122	.0200	76
1974	.163	.0250	76
1975	.093	.0096	76
1976	.041	.0058	78

Source: Based on Bureau of Labor Statistics WPI subfile tape.

*This column shows the total number of commodities for which the price indexes of year t, year t-1 and the corresponding weight, are all available.

Appendix

The procedure adapted in the calculation of the price dispersion series is described here in more detail.

1) Calculation of individual price rates of change.

The indexes used are annual averages. Their rates of change are computed by the first difference in the logarithms of the indexes, namely

$\Delta P_{it} = \log P_{it} - \log P_{it-1}$, where P_{it} and P_{it-1} are the price indexes of commodity i at time t and t-1.

2) Calculation of the weights assigned to each commodity.

Here I am refering briefly to the methodology of the computation of the WPI. For details see Wholesale Prices and Price Indexes (1973). The weights assigned to the commodities comprising the WPI are the percentages of the primary sales of the various commodities in the total of primary sales. The values of these weights are not available for each year. During the sample period, Industrial Censuses were carried on in 1952/53, 1954, 1958, 1963, and 1972. Major revisions of the weights were performed a few years after each census, in January of the years 1955, 1958, 1961, 1967, and 1976, when these data became available. From one year to the next between two major revisions, the weights were updated by the rate of change in the commodity price, and then normalized to get their sum equal to one. This procedure implies the assumption that the quantities are constant between two major revisions and that only the differential price changes affect the share of each commodity in the total sales.

Since I did not find available the computed weights for each year in the sample, I used the same procedure and the published weights data for December 1954, 1957, 1960, 1966, and 1975 to calculate the intermediate years figures. More concisely, having the weights for T = 1955, 1958, 1961, 1967, and 1976, the corresponding figures for the intermediate years are computed as follows

$$\tilde{w}_{it} = w_{it}\ (P_{it}/P_{iT})$$

$$w_{it} = \tilde{w}_{it}/\sum_i \tilde{w}_{it}$$

where w_{it} is the normalized weight and $T \leq t < T + 1$. For the period 1948-1954, T = 1955.

A likely minor additional point is that the weights assigned to each ΔP_{it} are the figures corresponding to the month of December in the previous year. Assuming that prices follow some trend, the annual average prices P_{it} and P_{it-1} would correspond to a pair of points somewhere in the middle of each one of the two concecutive years. Therefore it seems a good approximation to weight ΔP_{it} with weights corresponding to December t-1, which presumably lies between those two points.

3) The average and variance of the rates of price change are calculated according to

$$\Delta P_t = \sum_i w_{it}\ \Delta P_{it}\ , \text{ and}$$

$$\gamma_t^2 = \sum_i w_{it}\ (\Delta P_{it})^2 - (\Delta P_t)^2.$$

The number of commodities that were included each year in the calculations are those for which P_{it}, P_{it-1} and w_{it} are all available. The number of commodities for which the above holds is shown in Table I, column (3), along with the computed values of ΔP_t and γ_t^2.

FOOTNOTES

[1]Earlier, Mills (1927), and Graham (1930) observed similar correlations in the U.S. during the period 1920-1926, and in the German hyperinflation respectively.

[2] Related evidence was obtained by Cukierman and Wachtel (1978). They found a positive correlation between price dispersion and the variance of inflationary expectations across individuals.

[3]The FEDV variable. For its definition, see Barro (1978).

[4]However, Barro and Hercowitz (1979) found that this measure of unanticipated money is uncorrelated with the amounts of the revisions in the published money stock data, which can be considered as unperceived money at the time of the first publication.

[5]The output equation in Barro (1978) contains also a military conscription variable, which could also be considered as a real shock. However, the v_t variable values are theoretically considered unperceived as such during period t. The military conscription hardly can be thought of as being currently unperceived and therefore, it is not taken into account for v_t.

[6]This holds because $\alpha^d \lambda(z) = \alpha^d \frac{1}{\alpha^s(z) + \alpha^d} < 1$ for all z. Recall that $\alpha^s(z) > 0$ in all markets. Therefore also on average: $\alpha^d \lambda < 1$.

[7]The results are also unchanged if actual money growth, rather than its unexpected component, is used in the equation.

[8]The equation was also estimated using the sample 1948-1972. This regression showed a much poorer fit. There was no change in the effect of $(\tilde{m}_t - \tilde{m}_{t-1})^2$, but the coefficient of $(v_t - v_{t-1})^2$ was also statistically insignificant for this sample.

[9] However, if the inflation rate follows a random walk (plus a constant) as Parks specifies, the change in the inflation rate represents not only unexpected inflation (plus a constant), but also the change in expected inflation . Obviously, a rationalization of the observed correlation under the second interpretation, would be based on rather different considerations.

[10] Particularly, I was not able to find relative importance data based on the 1947 Census, which the BLS used as basic weights for computing the corresponding figures until December 1953. Therefore, the weights for the period 1948-1953 are constructed backwards using the data from the 1952-1953 Census.

[11] The relative importance data were obtained from Wholesale Prices and Price Indexes. (1958) p. 33, (1967), p. 73 and (1976) p.18.

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