NBER WORKING PAPER SERIES

INFLATION AND WAGE DISPERSION

Allan Drazen

Daniel S. Hamermesh

Working Paper No. 1811

NATIONAL BUREAU OF ECONOMIC RESEARCH
1050 Massachusetts Avenue
Cambridge, MA 02138
January 1986

The research reported here is part of the NBER's research program in Labor Studies. Any opinions expressed are those of the authors and not those of the National Bureau of Economic Research.

NBER Working Paper #1811
January 1986

Inflation and Wage Dispersion

ABSTRACT

A large body of empirical work has demonstrated that higher inflation, especially when it is unexpected, leads to greater dispersion in the distribution of price changes across subaggregates. A sparse and more recent literature suggests exactly the opposite effects on the distribution of wage changes. This study first reconciles these apparently opposite results using a model in which shocks to the economy can affect both wages and prices and the demand for indexing. If the positive effect of shocks on the demand for indexing is sufficiently large, the dispersion of changes in wages or prices will be reduced even though the shocks' direct effect is to increase this dispersion. Implicitly from the evidence, this offset is large enough in wage-setting, but not so large in price determination.

Additional evidence on the relationship between inflation and the dispersion of wage changes is provided by empirical work for 14 Israeli manufacturing industries, 1956-82. The results suggest that in Israel, just as in the United States (on which previous work has been conducted) with its much less rapid and variable inflation, dispersion also decreased with unexpected price inflation.

Allan Drazen
Department of Economics
Tel Aviv University
Ramat-aviv, ISRAEL

Daniel S. Hamermesh
Department of Economics
Michigan State University
East Lansing, MI 48824

I. Introduction

The relation between inflation and relative price variability has received increasing attention in recent years. Several authors have found that increased relative price variability is associated with increases in unanticipated inflation (Parks, 1978; Fischer, 1982; Cukierman and Wachtel, 1982, to name a few). The apparent pervasiveness of this phenomenon across countries (see, for example, Cukierman and Leiderman, 1981, for discussion of the Israeli experience) has led to a wide acceptance of the view that this positive correlation is (to use Friedman's terminology) an empirical "regularity" of economic systems. [1]

Hamermesh (1986) has considered the relation between unexpected inflation and relative _wage_ variability in the United States and found that the "regularity" does _not_ hold. Increases in unexpected inflation tended to _narrow_ the dispersion of wage changes across sectors in the period 1955-81. Allen (1984) finds similar results for the postwar U.S., but finds that unexpected inflation widened the dispersion of relative wage changes in the United States before World War II.

These results raise at least two questions. First, is the narrowing of wage dispersion in response to inflation unique to the postwar United States, or is it characteristic of other economies as well, particularly those with much higher and more variable rates of price inflation? Second, if lower wage variability is a general result, how might it be consistent with the above results on price variability? The purpose of this note is to address these two issues.

II. A General Framework

We begin by setting out a prototype model that can account for both a positive and a negative relation between inflation and price dispersion, depending on the nature of the price-setting process. The key question is whether prices are set before or after an aggregate nominal shock is observed. If price and quantity decisions must be made before the realization of the inflation shock is known, increased unanticipated inflation will increase relative price dispersion in a model where there is a confusion of aggregate and relative shocks. If, however, indexing arrangements allow prices to be adjusted after nominal shocks have been observed, a higher mean level of unanticipated inflation (that is, greater inflation uncertainty) may increase the degree of indexing and reduce price dispersion. In short, the effect of inflation on the dispersion of wage or price changes depends on the nature of price-setting--whether prices are set ex ante or ex post the realization of aggregate nominal shocks--and how inflation changes the nature of price-setting arrangements.

Let demand and supply for good i be log-linear of the form:

$$(1) \qquad y_t^d(i) = -\mu(i)\left(p_t(i) - p_t^*(i)\right) + \alpha\left(x_t - p_t^*(i)\right) + \omega_t(i),$$

and

$$(2) \qquad y_t^s(i) = \gamma(i)\left(p_t(i) - p_t^*(i)\right),$$

where $y_t^d(i)$, $y_t^s(i)$ are the logarithms of quantity demanded and supplied of good i, $p_t(i)$ is log price of good i, $p_t^*(i)$ is the perception by individuals in market i of the general price level, x_t is an aggregate nominal variable (so that the second right-hand side term in (1) can be thought of as a real-balance effect), and $\omega_t(i)$ is a random (excess) demand shock. $p_t(i) - p^*_t(i)$ is the locally perceived relative price, so that $\mu(i)$ and $\gamma(i)$ are demand and supply elasticities ($\mu, \gamma > 0$), which can vary across markets. This characteristic is crucial to the results.

The excess demand shock is assumed normal with mean zero and variance σ^2_ω, and $\omega_t(i)$ is uncorrelated over time and across markets. The rate of change of the nominal variable $x_t, \Delta x_t \equiv x_t - x_{t-1}$, obeys:

$$
\begin{aligned}
(3) \qquad \Delta x_t &= E_{t-1}(\Delta x_t) + \varepsilon_t \\
&= \delta_t + \varepsilon_t ,
\end{aligned}
$$

so that δ_t is that part of Δx_t which is predictable given information up to and including t-1 and is assumed known to individuals in all markets. The innovation ε_t is assumed normal with mean zero and variance σ^2_x, serially uncorrelated and independent of $\omega_t(i)$. One may note that $p_t(i)$ will in general convey information about ε_t and, since relative demand shocks differ across markets, how the posterior expectation of x_t varies across markets, leading to differential expectations about the current general price level $p^*_t(i)$. As a reference point we may then solve for the market-clearing price:

$$
(4) \qquad p_t(i) = \alpha\lambda_i[x_{t-1} + \delta_t] + [1 - \alpha\lambda_i]p^*_t(i) + \lambda_i[\alpha\varepsilon_t + \omega_t(i)],
$$

where $\lambda_i \equiv \dfrac{1}{\mu(i) + \gamma(i)}$. This general set-up is almost identical to that used in Cukierman and Leiderman (1984).

We model price-setting and indexing as follows. In each period the market-clearing price is observed, and the market-clearing quantity is transacted. Agents may decide that the price actually paid may be adjusted *ex post*, after ε_t has been observed. An indexing arrangement then takes the form:

$$(5) \qquad \hat{p}_t(i) = p_t(i) + \beta(i)\varepsilon_t$$
$$= \alpha\lambda_i[x_{t-1} + \delta_t] + [1 - \alpha\lambda_i]p_t^*(i) + \lambda_i[\alpha\varepsilon_t + \omega_t(i)] + \beta(i)\varepsilon_t ,$$

where $\beta(i)$ is an indexing parameter in the i'th market.

In the case where no indexing is used, $\hat{p}_t(i) = p_t(i)$. In the other polar case, where sector-specific demand shocks are not allowed to influence prices, (5) becomes:

$$(5') \qquad \hat{p}_t(i) = \alpha\lambda_i[x_{t-1} + \delta_t] + [1 - \alpha\lambda_i]p_t^*(i) + \beta(i)\varepsilon_t ,$$

where the first two terms in (5') are known at time t.

We can now compute the dispersion of equilibrium relative prices under different assumptions about the weights given to sector-specific demand shocks and *ex-post* indexing in setting individual prices. For simplicity, assume that the degree of indexing is the same across sectors, so that $\beta(i) = \beta$ for all i.

The general price level, p_t is a geometric weighted average of individual prices, namely:

$$(6) \qquad p_t = \sum_i u(i)\hat{p}_t(i); \qquad \sum_i u(i) = 1 ,$$

where $u(i)$ is the weight of the i'th good. p_t is a function of the $\hat{p}_t(i)$, which in turn depend on the expectation of p_t, namely $p_t^*(i)$. In equilibrium $p_t(i)$, p_t, and $p_t^*(i)$ are determined simultaneously. To find the solution, we use the method of undetermined coefficients. Since the model is log-linear, we hypothesize a solution:

$$(7) \qquad p_t = \pi_1 \delta_t + \pi_2 x_{t-1} + \pi_3 \varepsilon_t ,$$

where the π_i are to be determined.

The rational perception of p_t in market i is:

$$(8) \qquad \begin{aligned} p_t^*(i) &= E\left(p_t \middle| I_t(i)\right) \\ &= \pi_1 \delta_t + \pi_2 x_{t-1} + \pi_3 E\left(\varepsilon_t \middle| I_t(i)\right), \end{aligned}$$

where $I_t(i)$ is the information set in market i at t, which includes $p_t(i)$. (Remember we are assuming that the market-clearing price is known, even if it is not the transaction price). Since all other aggregate information (which is the same across markets) is independent of ε_t, it is information on $p_t(i)$ that is crucial in taking the expectation of ε_t. From (4), an observation on $p_t(i)$ is equivalent to an observation on $\alpha\varepsilon_t(i) + \omega_t(i)$. Given the normality of ε_t and $\omega_t(i)$, the optimal linear forecast is the least-squares projection of ε_t on $[\alpha\varepsilon_t + \omega_t(i)]$, yielding:

$$(9) \qquad E\left(\varepsilon_t \middle| I_t(i)\right) = \frac{\theta}{\alpha}\,[\alpha\varepsilon_t + \omega_t(i)], \text{ where } \theta = \alpha^2\sigma_x^2/[\alpha^2\sigma_x^2 + \sigma_\omega^2] .$$

We thus obtain:

$$(10) \qquad p^*_t(i) = \pi_1\delta_t + \pi_2 x_{t-1} + \pi_3 \frac{\theta}{\alpha}[\alpha\varepsilon_t + \omega_t(i)] \ ,$$

and

$$(11) \qquad \hat{p}_t(i) = \alpha\lambda_i[\delta_t + x_{t-1}] + [1 - \alpha\lambda_i]\left(\pi_i\delta_t + \pi_2 x_{t-1} + \pi_3 \frac{\theta}{\alpha}[\alpha\varepsilon_i + \omega_t(i)]\right) + \lambda_i[\alpha\varepsilon_t + \omega_t(i)] + \beta\varepsilon_t \ .$$

Substituting (11) into (6), rearranging, and assuming that the number of markets i is large, one obtains:

$$(12) \qquad \pi_1 = \pi_2 = 1; \qquad \pi_3 = \frac{\beta + \alpha\sigma}{1 - \theta + \alpha\sigma} \ ,$$

where $\sigma = \sum_i u(i)\lambda_i$.[2] .

Substituting into the expression for $\hat{p}_t(i)$ and rearranging terms:

$$(13) \qquad \hat{p}_t(i) = \delta_t + x_{t-1} + \frac{\lambda_i(1 - \theta[1 + \beta]) + \frac{\theta}{\alpha}[\beta + \alpha\sigma]}{1 - \theta + \alpha\theta\sigma}[\alpha\varepsilon_t + \omega_t(i)] + \beta\varepsilon_t \ .$$

Finally, we can compute p_t in equilibrium by substituting (13) into (6) and recognizing that the sum of terms in $\omega_t(i)$ converges (in probability) to zero:

$$(14) \qquad p_t = \delta_t + x_{t-1} + \left(\frac{\alpha\sigma + \beta}{1 - \theta + \alpha\theta\sigma}\right)\varepsilon_t \ .$$

We can now compute relative price variability as a function of unanticipated nominal shocks for a given degree of indexing. One measure of relative price variability is:

$$(15) \qquad S \equiv \underset{\omega}{E} \underset{i}{\Sigma} u(i)[\hat{p}_t(i) - p_t]^2 .$$

Using (13) and (14) in (15) yields:

$$S = \alpha^2(1 - \theta[1 + \beta])^2 \, DH(\lambda)\varepsilon_t^2$$

$$(16) \qquad + \left(\left\{(1 - \theta[1 + \beta])\sigma + \frac{\theta}{\alpha}[\beta + \alpha\sigma]\right\}^2 + (1 - \theta[1 + \beta])^2 H(\lambda)\right) D\sigma_\omega^2 ,$$

where $H(\lambda) = \sum_i u(i)(\lambda_i - \sigma)^2$ and $D = [1 - \theta + \alpha\theta\sigma]^{-2}$.

Equation (16) summarizes the determinants of relative price variability. The second term shows that relative price variability depends, naturally, on the varability of relative excess demand shocks. The first term shows that it depends on the level of the unanticipated nominal shock (an effect discussed by Hercowitz, 1981; 1982), with the size of the effect determined by the extent of indexing. Given β , a higher mean ε^2 induces higher price dispersion.

How will dispersion be related to β for a given realization of ε_t ? One immediately notices that for $\beta < \frac{1}{\theta} - 1$, increases in β will decrease the effect of given nominal shocks on dispersion. For increased aggregate nominal uncertainty (as measured by higher mean ε^2) to decrease relative price dispersion, it must therefore increase the indexing parameter β . That is, for a given β higher ε^2 clearly increases dispersion. If, however, higher ε^2 induces an increase in the degree of indexing β , dispersion can fall.

The source of an increase in β is not hard to find. If we consider the variance of a single relative price, we see that it too is related to β . Calling this variance S_i, we have:

$$(20) \quad S_i = \underset{\omega}{E}\,\underset{\varepsilon}{E}\left(p_t(i) - p_t\right)^2 ,$$

$$= \alpha^2\left(1 - \theta[1 + \beta]\right)^2[\lambda_i - \sigma_\varepsilon]^2 \cdot D\sigma_\varepsilon^2 + \lambda_i^2 \left\{\left(1 - \theta[1 + \beta]\right) + \frac{\theta}{\alpha}[\beta + \alpha\sigma]\right\}^2 \cdot D\sigma_\omega^2 .$$

As long as $\frac{\theta}{\alpha}$ is sufficiently small, an increase in β (for $\beta < \frac{1}{\theta} - 1$) will decrease S_i, while an increase in σ_ε^2 will increase S_i. Therefore, if individual utility is a decreasing function of the price variability of one price, an optimal response to an increase in σ_ε^2 is a higher degree of indexing.

This model can obviously be extended in a number of ways. Differential indexing would increase relative price (or wage) dispersion. More complicated indexing schemes (such as an explicit relation of the degree of indexing to observed economic variables) would enrich this model. Our purpose here was to present a model in which the effects of different methods of price-setting could be studied and the possibility of their different effects on the dispersion of relative prices and wages be explored.

III. Testing the Dispersion Hypothesis

The theoretical model suggests that, where indexing is a possibility, increases in unanticipated inflation may actually lower dispersion by increasing the degree of (implicit or explicit) indexing. As an evaluation of the applicability of these ideas, annual data from the Israeli economy were used to examine the model describing the variability of relative wage changes between 1956 and 1982. Data covering 14 manufacturing and mining industries, accounting for approximately 90 percent of output in these sectors, were used. [3/] With its rapid and highly variable inflation, Israel may be a good example of an economy in which wage-setting is characterized by changes in the degree of formal, and especially informal indexing as uncertainty about inflation varies.

Using the same equation for calculating the variance of relative wage changes as in Hamermesh (1986), we computed Var (w) as the weighted average of annual changes in the relative nominal daily wage rates of workers. [4/] The same formula was used to calculate the variance of changes in relative output, Var(y), with output being measured by the index of industrial production for each industry.

The annual inflation rate, p, was calculated as the twelve-month average of annualized monthly rates of change in the Consumer Price Index in Israel. Unlike the United States, in which several surveys on expectations about price inflation exist (and are used in Hamermesh, 1986), no comparable series exist in Israel. To test the model we thus construct inflation forecasts based on macro time series. The first forecast uses monthly rates of inflation to construct ARIMA forecasts for one, two, three, etc., up to twelve months ahead. In each case the most recent seven years of data on prices were used. Thus, for example, data on monthly inflation rates from 1948–1954 were used to construct an ARIMA forecast of the monthly inflation rates in 1955. While different ARIMA structures characterized different seven-year time periods in Israeli inflation in the past 35 years, in most cases integrated autoregressive processes involving one- and twelve-month lags had the greatest explanatory power. Thus the series p_a^e was constructed using these forecasts. Unexpected inflation using this forecast is $p - p_a^e$.

The second forecast of inflation is the naive projection of the previous year's rate of price inflation, p_{-1}. Such a simple forecast may prove adequate in an economy such as Israel's, in which inflation until the

1980's has seemed to fluctuate randomly on several plateaus between which there are discrete jumps. Unexpected inflation based on this forecast is $p - p_{-1}$.

The basic data (other than Var y) are shown in Table 1. It is apparent that, as in the United States, there is no particular trend in Var w. The series does show substantial variability, but some of the lowest values are observed in the early 1980s, the years of most rapid inflation in Israeli history. The forecasts p_a^e track the broad trends in inflation fairly well, but seem to predict poorly the yearly variation in inflation rates in the early part of the sample period. This is undoubtedly the result of the rapid annual fluctuations in inflation in the early 1950s and the lack of a significant trend term in the forecasting equations in that period.

The estimates of the determinants of Var w are presented in Table 2. The equations were estimated using the Cochrane-Orcutt adjustment for serial correlation in the errors. As in Hamermesh's (1986) results for the United States, increases in the dispersion of output shocks increase the dispersion of wage changes. As the estimates in the first column show, though, there is essentially no relation between Var w and the actual rate of price inflation. In both the United States and Israel, and contrary to the received wisdom about the effects of price inflation on the dispersion of relative prices, the data do not indicate any relation between inflation and relative wage dispersion.

As in the United States, however, once the series on price inflation is decomposed into expected and unexpected components we observe a striking relationship between inflation and Var w. Neither of the two inflation forecasts, p_{-1} or p_a^e , has a significant effect on the dispersion of relative wage changes. However, in both cases, and especially for the naive forecast,

Table 1. Variance of Annual Relative Wage Changes, Inflation Rate and ARIMA Forecast of Inflation, Israel, 1956-82

Year	Variance[a]	p	p^e_a
1956	.0860	4.059	17.19
1957	.0183	5.423	0.131
1958	.0537	4.363	-7.144
1959	.0512	2.334	-10.53
1960	.0519	3.560	1.228
1961	.0714	9.477	3.309
1962	.0374	10.28	7.670
1963	.0447	5.102	9.480
1964	.0207	4.582	7.246
1965	.0415	7.137	6.713
1966	.0162	7.861	8.274
1967	.0929	0.205	7.572
1968	.0818	1.959	.457
1969	.1033	3.925	0.842
1970	1.0726	10.22	3.317
1971	.0895	13.44	9.698
1972	.0950	12.42	13.77
1973	.1667	26.49	15.43
1974	.0690	26.49	31.92
1975	.0620	23.71	75.54
1976	.0951	38.24	39.62
1977	.5578	43.21	46.61
1978	.4611	48.34	48.90
1979	.2072	111.9	54.85
1980	.0981	112.7	107.7
1981	.0662	102.0	146.1
1982	.0563	131.8	147.7

[a]Actual variables x 10^2.

Table 2. Determinants of Var w, 14 Manufacturing Industries, Israel, 1956-82[a]

Constant	.00076	.00059	.00071
	(1.48)	(.87)	(1.25)
p	-.00028		
	(-.27)		
p_{-1}		.0010	
		(.76)	
$p-p_{-1}$		-.0067	
		(-3.09)	
p_a^e			-.00047
			(-.43)
$p-p_a^e$			-.0036
			(-1.67)
Var y	.079	.106	.090
	(4.91)	(6.86)	(5.47)
$\bar{R}^2$	.47	.60	.51
$\hat{\rho}$	.16	.46	.27
	(.73)	(2.21)	(1.22)

[a]t-statistics in parentheses.

p_{-1}, unexpected inflation, the deviation of the actual inflation rate from the forecast, has a negative effect on the dispersion of relative wage changes. The term representing unexpected inflation is highly significant in the equation using p_{-1}, and is significantly different from zero at the 90-percent level in the equation using p_a^e. 5/

Whether the very strong results using the forecast p_{-1} are more believable than those using the ARIMA forecast, p_a^e, depends partly on which forecast predicts annual inflation rates better. Some evidence on this question is presented in Table 3, showing regressions of the inflation rate on these forecasts. As the Table makes clear, the $\bar{R}^2$ is higher for the naive forecast than for the annualized monthly ARIMA forecast; and the latter does not add to the explanatory power of the lagged forecast when both are included in the same regression. We may conclude from this that the results based on the naive forecast deserve greater attention. 6/

IV. Conclusions

In this note we have shown how relative prices and wages are affected by shocks to relative demand and to the entire economy. We have demonstrated that their effects depend on whether they also affect the extent of indexing of price- and/or wage-setting. In particular, if greater absolute shocks to the economy increase the demand for indexation, those shocks can reduce dispersion in wage- and/or price-setting. Without any change in indexation, though, larger shocks will increase dispersion.

Earlier empirical work has shown that inflationary shocks have increased the dispersion of relative price changes in the postwar United States, while decreasing the dispersion of relative wage changes. Similarly, other work has demonstrated that (the much larger) inflationary shocks have

Table 3. Forecasts of Annual Inflation Rate, Israel, 1956-82

Constant	.024 (.72)	.047 (1.17)	2.41 (.71)
P_{-1}	1.10 (13.38)		1.02 (3.52)
P_a^e		.81 (10.49)	.07 (.29)
$\bar{R}^2$	.87	.80	.86

increased the dispersion of relative price changes in Israel. Our empirical work shows that, as in the United States, those shocks reduce the dispersion of relative wage changes in Israel. Taken as a whole and in conjunction with the model set forth here, the results indicate that inflationary shocks induce workers to seek (formal and informal) indexing to such an extent that relative-wage variability declines in response to the shock. While shocks may induce some increase in the indexing of price-setting arrangements, it is implicitly less extensive, as it is insufficient to offset the positive effects of the shocks on the observed dispersion in relative prices. The difference between wage- and price-setting in an environment of inflationary shocks may perhaps be due to the greater risk aversion of individuals in their roles as workers than as consumers.

REFERENCES

Steven Allen, "Changes in Relative Wages in the U.S., 1860-1983," Unpublished Paper, North Carolina State University, December 1984.

Alex Cukierman and Leonardo Leiderman, "Relative Price Variability and Inflation in the Israeli Economy," Discussion Paper 6-81, Pinchas Sapir Center for Development, Tel Aviv University, 1981.

____________, "Price Controls and the Variability of Relative Prices," Journal of Money, Credit and Banking, 16 (1984), 271-284.

____________ and Paul Wachtel, "Relative Price Variability and Nonuniform Inflationary Expectations," Journal of Political Economy, 90 (1982), 146-157.

Stanley Fischer, "Relative Price Variability and Inflation in the United States and Germany," European Economic Review, 18 (1982), 171-196.

Daniel Hamermesh, "Inflation and Labour-Market Adjustment," Economica, 53 (1986), forthcoming.

Zvi Hercowitz, "Money and the Dispersion of Relative Prices," Journal of Political Economy, 89 (1981), 328-356.

____________, "Money and Price Dispersion in the United States," Journal of Monetary Economics, 10 (1982), 25-37.

Richard Parks, "Inflation and Relative Price Variability," Journal of Political Economy, 86 (1978), 79-95.

FOOTNOTES

1/ One should note the finding of Hercowitz (1982). He argued that both relative price dispersion and unanticipated inflation may be thought of as endogenous, being driven by, among other things, monetary shocks. He found no significant correlation between unanticipated money shocks and relative price variability in postwar U.S. data.

2/ The intermediate step yields:

$$\begin{aligned} p_t &= \sum_i u(i)\left(\alpha\lambda_i + [1 - \alpha\lambda_i]\ \pi_1\right)\delta_t \\ &+ \sum_i u(i)\left(\alpha\lambda_i + [1 - \alpha\lambda_i]\ \pi_2\right)x_{t-1} \\ &+ \sum_i u(i)\left(\lambda_i + \frac{\theta}{\alpha}\ [1 - \alpha\lambda_i]\ \pi_3\right)\omega_t(i) \\ &+ \left(\beta + \alpha\sum_i u(i)[\lambda_i + \frac{\theta}{\alpha}\ [1 - \alpha\lambda_i]\pi_3]\right)\varepsilon_t\ , \end{aligned}$$

Its right-hand side equals that of (7), so that the coefficients must be equal. To obtain (12) we use the fact that, with a large number of "small" markets, the third term on the right side converges to zero in probability (when λ_i is bounded away from zero).

3/ The printing and publishing, diamond, basic metal, and miscellaneous manufacturing industries were excluded. For the former three data were not available for the entire period, while we felt that miscellaneous manufacturing was so heterogeneous that its composition would change frequently and thus induce errors into our estimates.

4/ The weights were the shares of industrial production accounted for by each industry. For 1955-1960 the 1955 weights were used; for 1961-1969 weights from 1963 were used; for 1970-1978 weights from 1970 were used, while 1979 weights were applied to the data from 1979-1982.

5/ It is also worth noting that the Box-Pierce test cannot reject the hypothesis that the residuals from this simple bivariate regression of p on p_{-1} are white noise. Using the contemporaneous correlation and three lags, and making Haugh's correction for degrees of freedom, the χ^2 test-statistic is 8.39. While this is significantly different from zero at the 90-percent level, it is not so at the 95-percent level.