STUART LANDON

University of Alberta
Edmonton, Alberta
Canada

Testing Aggregate Neutrality with Heterogeneous Sectors*

The existing empirical literature on the neutrality of aggregate demand policy implicitly assumes that the parameters of industry level output functions differ insignificantly across sectors. This paper tests this aggregation restriction and examines its implications for tests of aggregate neutrality. The results indicate that sectoral effects are significant determinants of aggregate output and that their exclusion may reverse the conclusions of aggregate neutrality tests.

1. Introduction

A large number of tests of the neutrality of aggregate demand policy, and of monetary policy in particular, have appeared in the literature.[1] These tests generally employ an estimating equation that relates aggregate output, unemployment or employment (or their growth rates) to, among other variables, the anticipated and unanticipated components of an aggregate demand variable—either the aggregate price level or the money supply. The unanticipated aggregate demand variable employed is generally proxied by the residual from an ad hoc forecasting equation. Aggregate demand shocks are said to be neutral if the coefficient on the anticipated aggregate demand variable is insignificantly different from zero. Various versions of this procedure have been criticized for employing inappropriate statistical tests and estimation strategies, for the inadequate specification of the forecasting

[1]See, for example, Sargent (1976), Barro (1977, 1978), Neftci and Sargent (1978), Fair (1979), Barro and Rush (1980), Leiderman (1980), Barro (1981), Hoffman and Schlagenhauf (1982), Mishkin (1982a, 1982b, 1983), Rush (1986), Gray, Kandil and Spencer (1987), and Chan (1988).

equation, and for the exclusion of important variables from the aggregate output and employment equations.[2]

Another potential shortcoming of the empirical neutrality literature is its reliance on the implicit assumption that the parameters of industry level output functions are identical (or differ insignificantly).[3] This assumption is necessary if the estimation of an aggregate output equation which does not explicitly incorporate diverse industry level responses is to yield unbiased parameter estimates and test results. Despite the importance and extreme nature of this aggregation restriction, it has not been tested and its impact on the conclusions of aggregate neutrality tests has not been examined.[4] The present paper analyzes this aggregation restriction and finds industry level effects to be important determinants of aggregate output. Their exclusion can yield biased parameter estimates and can reverse the conclusions of neutrality tests.

The paper is divided into five sections. An aggregate output equation incorporating different industry level output responses is derived in Section 2. Section 3 describes the data and estimation methodology. In Section 4 the estimation results are examined, the significance of different industry level effects is investigated, and the neutrality hypothesis is tested. Some brief concluding comments are provided in Section 5.

2. Derivation of the Aggregate Output Function

Following Lucas (1973), the log of real output in sector z during period t is given by:

$$y_{zt} = y_{zt}^n + \phi_z[P_{zt} - E(P_t|I_t(z))] + \beta_z y_{z,t-1}, \qquad z = 1, \ldots, m, \qquad (1)$$

where

y_{zt}^n = the log of the natural level of sector z output;

P_{zt} = the log of the price of sector z output;

$E(P_t|I_t(z))$ = the expected value of the log of the aggregate price level conditioned on information available in sector z at time t. This information consists of the values of all $t-1$ period variables as well as the period t price of the good produced in sector z;

---

[2] Critiques of the empirical aggregate neutrality literature can be found in, for example, Pesaran (1982), Mishkin (1983) and Pagan (1984).

[3] This assumption goes back to Lucas (1973).

[4] Kretzmer (1989) estimates industry level supply equations but does not examine the impact of diverse industry level behavior on aggregate output, nor does he test neutrality. Gauger and Enders (1989) find that unanticipated aggregate shocks alter the composition of output, but not aggregate output.

$y_{z,t-1}$ = the log of sector z output in period $t-1$;
ϕ_z, β_z = constant parameters.

Equation (1) implies that the principal determinant of output fluctuations in a particular sector is variations in the price of that sector's output relative to the expected *aggregate* price level.

   The values of all $t-1$ period variables are known economy wide in period t and can be used to determine a distribution for P_t that is common to all agents and has a mean and variance given by $\overline{P}_t$ and σ_P^2 respectively. The level of the price of the good produced in sector z differs from P_t by an amount v_{zt} which is independent of P_t and has mean zero and variance σ_z^2. That is,

$$P_{zt} = P_t + v_{zt} . \tag{2}$$

   In period t, agents in sector z know P_{zt} and the distributions, but not the values, of P_t and v_{zt}. Using this information and Equation (2), it is possible to determine the expected value of the aggregate price level which can then be substituted into Equation (1) to yield

$$y_{zt} = y_{zt}^n + \alpha_z[P_{zt} - \overline{P}_t] + \beta_z y_{z,t-1} , \tag{3}$$

where $\alpha_z = \phi_z \Omega_z$ and $\Omega_z = (\sigma_z^2)/(\sigma_P^2 + \sigma_z^2)$.

   Average output across all m sectors, y_t, is given by

$$y_t = \sum_z \theta_{zt} y_{zt} , \tag{4}$$

where θ_{zt} is the proportion of total output produced in sector z during period $t-1$.[5] Combining Equations (3) and (4) yields

$$y_t = y_t^n + \sum_z \theta_{zt}\alpha_z[P_{zt} - \overline{P}_t] + \sum_z \theta_{zt}\beta_z y_{z,t-1} , \tag{5}$$

where $y_t^n = \sum_z \theta_{zt} y_{zt}^n$. This is equivalent to

$$y_t = y_t^n + \alpha_m[P_t - \overline{P}_t] + \beta_m y_{t-1}$$

$$+ \sum_{z=1}^{m-1} (\alpha_z - \alpha_m)\theta_{zt}[P_{zt} - \overline{P}_t] + \sum_{z=1}^{m-1} (\beta_z - \beta_m)\theta_{zt}y_{z,t-1} , \tag{6}$$

where $P_t = \sum_z \theta_{zt}P_{zt}$, $y_{t-1} = \sum_z \theta_{zt}y_{z,t-1}$ and α_m is the α_z parameter for sector m. The last two terms in Equation (6) represent the impact on aggregate

[5]The use of proportions from the previous period implies that the weights used to calculate average output are independent of the current distribution of output across sectors. Use of period t proportions alters the results trivially.

output of diverse reactions at the industry level to anticipated relative price changes and different rates of growth.[6]

If α_z and β_z are the same for all m sectors, Equation (6) can be rewritten as

$$y_t = y_t^n + \alpha[P_t - \overline{P}_t] + \beta y_{t-1}, \tag{7}$$

where $\alpha = \alpha_z$ and $\beta = \beta_z$ for all $z = 1, \ldots, m$. Equation (7) is the output equation which forms the basis for most tests of the neutrality hypothesis. The assumption that α_z and β_z are the same for all sectors has two important implications. First, identical β_z parameters imply that the speed of adjustment in each sector is the same and that the distribution of output across sectors has no effect on the rate of change of aggregate output. Second, if the α_z coefficients are identical, a shift in the distribution of industry level prices which leaves P_t unchanged, no matter how extreme, has no impact on aggregate output.

Two conditions are necessary for α_z to be identical across all m sectors. First, the parameter representing the supply response to relative price changes, ϕ_z in Equation (1), must be the same for all z. Second, the variance of the deviation of the price level in sector z from the economy wide average (σ_z^2) must be identical for all m sectors. This second condition implies that the quality of information on aggregate prices reflected in each sector's own price must be the same in every sector.

Imposing identical α_z and β_z coefficients on the data may be overly restrictive. Industries are characterized by different technologies, face different input supply functions, and must bear different adjustment costs.[7] If these factors cause α_z and β_z to differ significantly across sectors, the specification given in Equation (7), by imposing zero restrictions on important sectoral variables, will cause the estimates of α and β and their standard errors to be biased (Theil 1971). This bias will extend to the estimated coefficients and standard errors of actual or anticipated aggregate demand variables if they are included in this equation for the purpose of testing neutrality.

The assumption that α_z and β_z do not differ significantly across sectors (and the $2(m-1)$ restrictions it implies) can be tested using estimates of Equations (6) and (7). These estimates can also be used to examine the sensitivity of aggregate neutrality tests to the imposition of identical industry level α_z and β_z parameters.

---

[6]Note that the same weights are used to calculate y_t, y_{t-1} and P_t and that these weights evolve through time.

[7]Ahmed (1987) provides empirical evidence that the α_z parameters differ significantly across industries.

3. The Data and the Estimation Methodology

Equations (6) and (7) are estimated using annual U.S. data from 1950 to 1989 for seven sectors: construction (CON); mining (MIN); manufacturing (MAN); retail and wholesale trade (TRA); finance, insurance, real estate and services (FIS); transportation, communication, electric, gas and sanitary services (TCE); and agriculture (AGR).[8] (Sources for this data as well as for the other data used are given in the Appendix.) As can be seen from Equation (6), data for only six of these seven sectors enter the estimating equation as independent explanatory variables.[9]

Tests for a unit root in the log levels of the industry price and output series, as well as the weighted averages of these series, could not reject the hypothesis of a unit root in all cases.[10] As a result, Equations (6) and (7) were estimated in first difference form.[11] (The first difference of a variable is denoted by the addition of a D prefix.)

As is standard in the literature, the expected aggregate price variable, $\overline{DP}_t$, is proxied by the predicted value from a DP_t forecasting equation. This equation should be parsimonious, but not exclude information that could significantly improve its forecasting ability. It seems reasonable that, at a minimum, it include an aggregate demand proxy, a monetary variable, and a variable representing supply-side shocks. On this basis, the independent variables in the forecasting equation, all of which are observable at the end of the period prior to that being forecast, were chosen to be the growth rate of lagged aggregate output (Dy_{t-1}), the growth rate of the money stock (DM_{t-1}), the growth rate of an energy price index (DPE_{t-1}) and DP_{t-1}.[12] Initially the price forecasting equation was estimated with two lags of each

[8]Annual data is employed because data for nominal and constant dollar GNP by industry is not available at more frequent intervals.

[9]The results are invariant to which sector is not included as an independent variable in this equation.

[10]These tests were conducted using the methodology outlined in Fuller (1976) and Dickey, Bell and Miller (1986). The unit root test equation took the form: $\Delta y_t = \mu + \alpha T + \theta y_{t-1} + \varepsilon_t$, where y_t is the variable being tested for a unit root, Δ is the difference operator, T is a time trend, μ, α and θ are constant parameters, and ε_t is a white noise error. The hypothesis of a unit root is tested by comparing the t-statistic associated with the estimated value of θ with the appropriate critical value in Fuller (1976). If the tests are carried out with the α parameter restricted to zero, a unit root cannot be rejected in 15 of the 16 series with the one rejection being for an individual price series.

[11]When taking first differences, the product of the change in the sectoral weights $(\theta_{zt} - \theta_{zt-1})$ and the difference between the corresponding sector's lagged price and the lagged anticipated price is set equal to zero for all z. These products are second-order small and their exclusion facilitates estimation.

[12]These variables were chosen prior to commencing estimation of the output equation and were not changed thereafter. DP_{t-1} is the difference between two price indices which are calculated using the sectoral weights from the same period as their component prices.

explanatory variable, but the coefficients on the twice-lagged variables were neither individually nor jointly significant and thus they were excluded from all further estimation.

The change in the natural level of aggregate output, Dy_t^n, is proxied by a constant term and the change in the standard deviation of industry level employment growth rates (DSIG).[13] This last variable is defined as in Lilien (1982) and is intended to proxy for changes in the natural rate which can be attributed to the slow movement of workers from declining to expanding sectors. The principal reason for including this variable is to determine whether the sectoral effects examined here can account for the significance of SIG as found by Lilien and, in addition, whether SIG can proxy for the sectoral aggregation effects introduced here.[14]

Combining the aggregate output equation derived in Section 2 with the assumptions made above yields the two estimating equations:[15]

$$DP = \lambda_0 + \lambda_1 DP_{-1} + \lambda_2 Dy_{-1} + \lambda_3 DM_{-1} + \lambda_4 DPE_{-1} + \epsilon_P, \tag{8}$$

$$
\begin{aligned}
Dy = {}& \gamma_0 + \gamma_1 DSIG + \gamma_2 \overline{DP} + \alpha_{AGR}(DP - \overline{DP}) + \beta_{AGR} Dy_{-1} \\
& + (\alpha_{CON} - \alpha_{AGR})\theta_{CON}(DP_{CON} - \overline{DP}) + (\alpha_{TCE} - \alpha_{AGR})\theta_{TCE}(DP_{TCE} - \overline{DP}) \\
& + (\alpha_{TRA} - \alpha_{AGR})\theta_{TRA}(DP_{TRA} - \overline{DP}) + (\alpha_{FIS} - \alpha_{AGR})\theta_{FIS}(DP_{FIS} - \overline{DP}) \\
& + (\alpha_{MAN} - \alpha_{AGR})\theta_{MAN}(DP_{MAN} - \overline{DP}) + (\alpha_{MIN} - \alpha_{AGR})\theta_{MIN}(DP_{MIN} - \overline{DP}) \\
& + (\beta_{CON} - \beta_{AGR})\,(\theta_{CON}y_{CON,-1} - \theta_{CON,-1}y_{CON,-2}) \\
& + (\beta_{TCE} - \beta_{AGR})\,(\theta_{TCE}y_{TCE,-1} - \theta_{TCE,-1}y_{TCE,-2}) \\
& + (\beta_{TRA} - \beta_{AGR})\,(\theta_{TRA}y_{TRA,-1} - \theta_{TRA,-1}y_{TRA,-2}) \\
& + (\beta_{FIS} - \beta_{AGR})\,(\theta_{FIS}y_{FIS,-1} - \theta_{FIS,-1}y_{FIS,-2}) \\
& + (\beta_{MAN} - \beta_{AGR})\,(\theta_{MAN}y_{MAN,-1} - \theta_{MAN,-1}y_{MAN,-2}) \\
& + (\beta_{MIN} - \beta_{AGR})\,(\theta_{MIN}y_{MIN,-1} - \theta_{MIN,-1}y_{MIN,-2}) + \epsilon_y, \tag{9}
\end{aligned}
$$

[13]This specification was chosen prior to commencing estimation and was not changed thereafter.

[14]The use of SIG has been criticized by Abraham and Katz (1986). The results with respect to the significance of the sectoral effects and the tests of neutrality are similar to those reported below, even if SIG is excluded from the model. If SIG is filtered of demand effects by regressing each industry employment growth rate (relative to total employment growth) on the anticipated aggregate price level and the relevant sector's price (or, alternatively, on anticipated and unanticipated aggregate prices) and then forming SIG using the residuals from these regressions, the test statistics associated with the zero restrictions on the sectoral variables all become slightly larger than those in Table 3. Thus, the conclusions with respect to the significance of the sectoral variables remain unchanged. The pattern of significance of the aggregate anticipated price level is also unaffected by changing the method of calculating SIG.

[15]Much of the empirical literature uses money surprises rather than the price surprises used here. The use of price surprises implies that it is not necessary to assume an immediate and stable relationship between money and prices (see Chan 1988) or to determine the direction and extent to which the aggregate money supply affects the prices of individual sectors. Price surprises are used in Sargent (1976), Fair (1979), Mishkin (1982a), and Gray, Kandil and Spencer (1987).

where $\overline{DP}$ is the predicted value from Equation (8), ϵ_P and ϵ_Y are independently distributed random errors with mean zero, and the time subscripts have been dropped for convenience.[16] The α_z, β_z and θ_z parameters in Equation (9) correspond to the same parameters in Equation (6). The γ_0 and $\gamma_1 DSIG$ terms proxy the natural level of output growth while $\gamma_2 \overline{DP}$ is introduced in order to test aggregate neutrality.[17]

A test of the hypothesis that industry level price responses are identical across sectors involves determining whether the six $(\alpha_z - \alpha_{AGR})$ parameters are jointly insignificant (where z denotes the six industries other than agriculture). A test of the hypothesis that output adjustment takes place at the same rate in all industries requires testing the joint significance of the six $(\beta_z - \beta_{AGR})$ parameters, while the neutrality of aggregate demand policy can be tested by determining whether γ_2 is significantly different from zero.

As noted in Pagan (1984,1986), for many types of models that include generated regressors, standard two-step estimation techniques (as used in Barro 1977, for example) can yield estimates of the parameters or standard errors which are inconsistent or inefficient. For this reason, the parameters of Equations (8) and (9) were estimated using the consistent and asymptotically efficient estimation method described in Proposition 3.5 of Pagan (1986). This estimation methodology involves initially estimating the parameters of the model using a two-stage estimation procedure and then updating these estimates by regressing the normalized estimated residuals from the initial procedure on the derivatives of the model (evaluated at the initial parameter estimates) in a double-length regression.

The estimation procedure suggested by Pagan (1986), and described in the previous paragraph, relies on aymptotic distribution theory. However, only 40 annual observations are available to estimate the 22 parameters which appear in the most unrestricted version of the two-equation model. One way to approximate the small sample distributions of the estimated parameters and test statistics is to employ a Monte Carlo procedure to bootstrap the

---

[16]As noted by Mishkin (1982b), Abel and Mishkin (1983) and Hoffman, Low and Schlagenhauf (1984), identification of the γ_2 parameter in Equation (9) requires that the covariance of ϵ_P and ϵ_Y be zero. This assumption is standard in the empirical work in this area and arises because the estimated value of ϵ_P is a regressor in the Dy_t equation. Hoffman, Low and Schlagenhauf (1984), using Monte Carlo experiments, find that test results are not significantly biased if this restriction is imposed and the covariance is not actually zero. Abel and Mishkin (1983) show that tests are unaffected by this restriction if there are no unanticipated prices as lagged regressors. This coincides with the case here due to the evidence given in footnote 17 below.

[17]Equation (9) does not include lagged price variables because the hypothesis that the coefficients on these variables are all zero could not be rejected using a 95% confidence interval. The likelihood ratio statistic for this test is 9.10 and the critical value is 14.07.

results. All the parameter estimates and test statistics reported in Tables 1 and 2 have been estimated using a bootstrap procedure.[18]

4. The Results

Table 1 gives the estimated parameters for four different versions of Equations (8) and (9).[19] These are, respectively: the unrestricted case, the case with neutrality imposed (γ_2 set equal to zero), the case with identical lagged adjustment effects (β_z equal to β_{AGR} for all z), and the case with identical industry level responses to anticipated relative price shocks (α_z equal to α_{AGR} for all z). Table 2 gives estimates of the parameters of Equations (8) and (9) when identical α_z and β_z parameters are assumed for all seven sectors (with and without aggregate neutrality imposed).

A comparison of the results in Tables 1 and 2 shows that the models which allow for different sectoral effects explain considerably more of the variance of the growth rate of aggregate output than do the models which assume identical sectoral effects. Seventeen of the 28 α and β parameters in the first two columns of Table 1 are significant at the 10% level.[20] When identical α_z and β_z parameters are imposed across all seven sectors, as in

[18]The bootstrap procedure used here involves first estimating the model using the Pagan (1986) methodology and calculating the estimated residuals for each equation. Assigning equal probability to each estimated residual and drawing from the sample of estimated residuals with replacement, 1000 new 40 observation samples of residuals were created for each equation. In conjunction with the initial estimates of the parameters and the exogenous right-hand-side variables, these 1000 new samples of residuals were used to create 1000 new series for each of the two dependent variables in the model. Using these series for the dependent variables and employing the Pagan methodology once again, the model was re-estimated 1000 times. For each of these 1000 replications all the parameters of the model and test statistics were calculated. The resulting 1000 values for each parameter and test statistic were then used to generate distributions for these same parameters and test statistics. The parameters and test statistics reported in Tables 1 and 2 are the means of these distributions while the reported standard errors of the parameters are the standard errors of the bootstrap distributions. Confidence intervals for each parameter are calculated using the percentile method since the histograms of the parameter distributions appear to be relatively symmetric. For examples and descriptions of the bootstrap methodology employed here see Freedman and Peters (1984a, 1984b), Efron and Tibshirani (1986) and Davidson and Mackinnon (1993). The major conclusions reported below are unchanged if the parameter estimates and test statistics are not bootstrapped.

[19]All versions of the model were estimated with rationality imposed. Rationality was tested and could not be rejected for the version of the model which did not include any sectoral effects. Similar tests could not be carried out with the sectoral model because of insufficient degrees of freedom.

[20]Note that the results in Table 1 imply that some of the underlying α_z parameters (as opposed to the $(\alpha_z - \alpha_{AGR})$ parameters) are negative. For example, this is the case for two (α_{MAN} and α_{MIN}) of the seven estimated α_z parameters in column I of Table 1. However, neither of these negative parameters is statistically different from zero (using a 95% confidence interval).

TABLE 1. *Estimates of the Parameters of Equations (8) and (9) When the Sector Specific Parameters are Allowed to Differ*

Coefficient	Unrestricted Case	Neutrality Imposed	Identical βs across Sectors	Identical αs across Sectors
	I	II	III	IV
λ_0	0.0003	−0.0005	−0.0006	0.0016
	(0.0055)	(0.0054)	(0.0054)	(0.0055)
λ_1	0.6260*	0.6313*	0.6461*	0.5698*
	(0.1350)	(0.1351)	(0.1271)	(0.1378)
λ_2	0.2435*	0.2474*	0.2465*	0.2442*
	(0.0739)	(0.0736)	(0.0757)	(0.0716)
λ_3	0.0001	0.0001	0.0001	0.0001
	(0.0001)	(0.0001)	(0.0001)	(0.0001)
λ_4	0.0728*	0.0767*	0.0686*	0.0881*
	(0.0343)	(0.0354)	(0.0308)	(0.0347)
γ_0	0.0328**	0.0243**	0.0420*	0.0324*
	(0.0166)	(0.0145)	(0.0100)	(0.0160)
γ_1	−1.018*	−1.063*	−0.9642*	−1.130*
	(0.2356)	(0.2446)	(0.2544)	(0.2973)
γ_2	−0.2389		−0.3889*	−0.3611*
	(0.1835)		(0.1880)	(0.1637)
α_{AGR}	−0.0117	−0.0007	−0.0117*	−0.0102
	(0.2506)	(0.2620)	(0.2721)	(0.2771)
β_{AGR}	−1.422**	−1.534*	0.1688	−0.3807
	(0.7489)	(0.7164)	(0.1368)	(0.6190)
$\alpha_{CON} - \alpha_{AGR}$	3.662*	3.565*	2.996*	
	(1.235)	(1.231)	(1.143)	
$\alpha_{TCE} - \alpha_{AGR}$	0.5679	−0.3563	2.207	
	(2.132)	(2.003)	(2.013)	
$\alpha_{TRA} - \alpha_{AGR}$	0.2468	0.3098	−0.2397	
	(0.6405)	(0.6832)	(0.6498)	
$\alpha_{FIS} - \alpha_{AGR}$	1.213	1.454**	0.7766	
	(0.7890)	(0.7702)	(0.7478)	
$\alpha_{MAN} - \alpha_{AGR}$	−1.403**	−1.063	−0.3488	
	(0.8025)	(0.7615)	(0.6955)	
$\alpha_{MIN} - \alpha_{AGR}$	−0.6742	−1.080**	−0.3461	
	(0.6415)	(0.5737)	(0.6343)	
$\beta_{CON} - \beta_{AGR}$	1.265*	1.437*		0.7148
	(0.4780)	(0.4615)		(0.4581)

(continued)

TABLE 1. *(continued)*

	Unrestricted Case	Neutrality Imposed	Identical βs across Sectors	Identical αs across Sectors
$\beta_{TCE} - \beta_{AGR}$	0.5167	0.5941		0.3502
	(0.3606)	(0.3765)		(0.3921)
$\beta_{TRA} - \beta_{AGR}$	1.365*	1.447*		0.7522**
	(0.4613)	(0.4452)		(0.4374)
$\beta_{FIS} - \beta_{AGR}$	1.670*	1.748*		0.9502*
	(0.4399)	(0.4335)		(0.4207)
$\beta_{MAN} - \beta_{AGR}$	1.414*	1.500*		0.7189
	(0.4875)	(0.4715)		(0.4417)
$\beta_{MIN} - \beta_{AGR}$	1.428*	1.420*		0.3748
	(0.7099)	(0.6845)		(0.5720)
Log of the Likelihood	270.51	268.68	259.22	260.40
$R^2 - Dy$	0.793	0.776	0.634	0.652
$- DP$	0.760	0.767	0.763	0.764
Godfrey test for AR1:				
$- Dy$	1.66†	1.48†	1.24†	1.47†
$- DP$	1.36†	1.44†	1.52†	1.24†
Breusch-Pagan test for heteroscedasticity (degrees of freedom):				
$- Dy$	15.01 (17)††	13.91 (16)††	10.30 (11)††	9.07 (11)††
$- DP$	3.49 (5)††	3.60 (5)††	3.66 (5)††	3.62 (5)††
Engle test for ARCH:				
$- Dy$	0.715†††	0.754†††	0.846†††	0.823†††
$- DP$	0.856†††	0.807†††	0.848†††	0.859†††

NOTES: The numbers in brackets under each estimated coefficient are the bootstrapped standard errors. The R-squared is calculated as the total sum of squares minus the sum of squared residuals all divided by the total sum of squares. Godfrey's test for AR1 is distributed as a χ^2 (1) as is the Engle test for autoregressive conditional heteroscedasticity. The Breusch-Pagan test is distributed as a χ^2 with the degrees of freedom given in brackets following the test statistic. Since the output equation is non-linear, the right-hand-side variables in the Godfrey and Breusch-Pagan tests are the partial derivatives of the DY equation with respect to each parameter.

*Significantly different from zero using the bootstrapped percentile method 95% confidence interval.

**Significantly different from zero using the bootstrapped percentile method 90% confidence interval.

†Cannot reject the null hypothesis of no serial correlation at a 95% confidence level.

††Cannot reject the null hypothesis of homoscedasticity using a 95% confidence level.

†††Cannot reject the null hypothesis of no autoregressive conditional heteroscedasticity at a 95% confidence level.

TABLE 2. *Estimates of the Parameters of Equations (8) and (9) When the Sector Specific Parameters are Not Allowed to Differ*

	Unrestricted Case	Neutrality Imposed
Coefficients	I	II
λ_0	0.0015	−0.0009
	(0.0057)	(0.0056)
λ_1	0.5915*	0.5976*
	(0.1357)	(0.1432)
λ_2	0.2417*	0.2563*
	(0.0743)	(0.0723)
λ_3	0.0001	0.0002
	(0.0001)	(0.0001)
λ_4	0.0807*	0.0951*
	(0.0344)	(0.0367)
γ_0	0.0489*	0.0307*
	(0.0084)	(0.0058)
γ_1	−1.000*	−1.229*
	(0.2828)	(0.2877)
γ_2	−0.4544*	
	(0.1540)	
α	−0.0089	0.0033
	(0.2726)	(0.2958)
β	0.1938	0.2005
	(0.1154)	(0.1237)
Log of the Likelihood	252.50	248.61
$R^2 - Dy$	0.484	0.352
$- DP$	0.759	0.787
Godfrey test for AR1:		
$- Dy$	1.08[†]	1.07[†]
$- DP$	1.37[†]	1.28[†]
Breusch-Pagan test for heteroscedasticity (degrees of freedom):		
$- Dy$	3.82 (5)[††]	2.98 (4)[††]
$- DP$	3.49 (5)[††]	3.57 (5)[††]
Engle test for ARCH:		
$- Dy$	0.880[†††]	0.772[†††]
$- DP$	0.778[†††]	0.807[†††]

See the notes to Table 1.

TABLE 3. *Tests for the Significance of Sectoral Effects*

	Unrestricted Case	Neutrality Imposed
Test of the hypothesis that α_z and β_z are identical across all seven sectors.	36.02*	40.14*
Test of the hypothesis that the α_z are identical across all seven sectors when the β_z can differ.	15.80*	17.44*
Test of the hypothesis that the β_z are identical across all seven sectors when the α_z can differ.	13.44*	16.40*

NOTE: All tests are likelihood ratio tests. The degrees of freedom of the three tests are, respectively, 12, 6 and 6. At a 95% confidence level the corresponding χ^2 critical values are: 21.03, 12.59, 12.59.
    *Rejects the null hypothesis at the 95% confidence level.

Table 2, neither the estimated value of α, nor that of β, is significant using a 90% confidence interval.

Several diagnostic statistics can be used to examine the robustness of the estimated results. A test for serial correlation, due to Godfrey (1978), cannot reject the hypothesis of no serial correlation in every case. Similarly, a Breusch-Pagan (1979) test for heteroscedasticity, which is also a general test for misspecification, cannot reject the hypothesis of homoscedasticity in every case. Finally, the test for autoregressive conditional heteroscedasticity of Engle (1982) finds no evidence of this form of heteroscedasticity.[21]

Table 3 provides likelihood ratio statistics for several different tests of whether the α_z and β_z parameters are identical across all seven sectors (with and without neutrality imposed). It can be seen from these test statistics that the hypothesis that all industry level α_z and β_z parameters are identical is rejected whether or not neutrality is imposed. A similar result follows for the test of whether the α_z parameters are identical when the β_z parameters are allowed to differ across sectors as well as for the test of the hypothesis that the β_z parameters are identical when the α_z parameters are allowed to vary.

[21] If the two equations of the model are assumed to be trend stationary, rather than difference stationary, the estimates are characterized by extremely high R^2 values as well as considerable evidence of serial correlation in the price equation and in the output equation when it does not include any sectoral effects. As Granger and Newbold (1974) point out, results of this type are indicators of a spurious regression. In conjunction with the results of the unit root tests reported above, these results indicate that the assumption of trend stationarity is unlikely to be an appropriate empirical methodology in this case.

As can be seen from column I of Table 1, the aggregate neutrality hypothesis cannot be rejected if α_z and β_z are allowed to vary across sectors (that is, γ_2 is not significantly different from zero in column I of Table 1).[22] However, if α_z is allowed to vary, but the β_z parameters are assumed to be identical (column III of Table 1) or if the α_z parameters are assumed to be equal across all seven sectors (column IV of Table 1), the neutrality hypothesis is rejected. Similarly, if both the α_z and β_z parameters are assumed to be equal across all seven sectors (column I of Table 2), the neutrality hypothesis is also rejected. In these latter three cases, anticipated aggregate price shocks have a significant negative impact on aggregate output.[23] These results imply that the rejection of neutrality is strongly dependent upon the imposition of identical sectoral effects. This could be one reason why previous empirical studies, none of which explicitly allows for different industry level responses, have found evidence that seems to contradict the neutrality hypothesis.[24]

The issue of the impact on aggregate output of industry level variables was raised by Lilien (1982). He showed that the standard deviation of industry level employment growth rates (SIG), a proxy for changes in the natural rate due to the slow movement of workers from declining to expanding industries, is an important determinant of unemployment. It is possible that the significance of the SIG variable used by Lilien follows from it acting as a proxy for the cross-industry aggregation affects addressed here. However, the evidence provided in Tables 1 and 2 suggests that SIG and the industry level effects included in Table 1 represent different factors determining aggregate output. The parameter associated with SIG, γ_1, is significant in all estimated

---

[22]Even when allowing the extent of non-neutrality to vary across sectors, a test of the neutrality hypothesis, that is, that the sectoral non-neutrality effects are all zero, cannot be rejected. The χ^2 statistic for this test is 10.63 while the critical value for the test using a 95% confidence interval is 14.07.

[23]The value of the significant negative coefficient associated with $\overline{DP}$ in Table 2, column I, is determined by the omitted variable bias which follows from excluding the sectoral price and lagged output variables from this equation. This omitted variable bias is given by the sum of the true coefficients on the excluded variables in the aggregate output equation each pre-multiplied by the coefficient associated with $\overline{DP}$ in a regression of each excluded variable on all the variables included in the model of column I in Table 2 (see Theil 1971). The negative bias in the estimate of γ_2 is due to this sum being dominated by excluded variables which have correlations with aggregate output and $\overline{DP}$ which are opposite in sign. This is the case for $(DP_{MIN} - \overline{DP})$ as well as for the lagged output variables of the construction and trade sectors. Both of these lagged output variables are negatively related to anticipated inflation, but have positive coefficients in the aggregate output equation. The rate of price increase in the mining sector is positively related to anticipated aggregate inflation, but increases in the mining sector's price relative to the aggregate anticipated price has a negative impact on aggregate output.

[24]See, for example, Mishkin (1982a, 1982b), Hoffman and Schlagenhauf (1982) and Chan (1988).

versions of the model and its value is relatively unaffected by the inclusion or exclusion of the other industry level variables from the aggregate output equation. Furthermore, as noted above, even with SIG included in the estimating equation, the industry level variables are significant determinants of aggregate output. As a result, SIG is unlikely to be a good proxy for the sectoral aggregation effects examined here.[25]

5. Conclusion

This paper tests the restriction, implicit in the empirical literature on aggregate neutrality, that industry level price and lagged adjustment responses differ insignificantly across sectors. If this restriction does not hold, estimates of the parameters (and corresponding standard errors) of an aggregate output equation which does not allow for different industry level responses will be biased.[26] The results provided above indicate that industry level effects are significant determinants of aggregate output. If these sectoral effects are arbitrarily neglected, the results indicate that conclusions with respect to the rejection of aggregate neutrality may be incorrect. While these results follow from the estimation of a relatively simple model for a particular historical period, they illustrate the potential for the imposition of invalid aggregation restrictions to yield biased parameter estimates and test results in aggregate equations.

Received: March 1993
Final version: April 1994

References

Abel, Andrew, and Frederic Mishkin. "An Integrated View of Tests of Rationality, Market Efficiency and the Short-Run Neutrality of Monetary Policy." *Journal of Monetary Economics* 11 (1983): 3–24.

Abraham, Katherine G., and Lawrence F. Katz. "Cyclical Unemployment: Sectoral Shifts or Aggregate Disturbances?" *Journal of Political Economy* 94 (1986): 507–22.

[25]If SIG is filtered of demand effects, as discussed in footnote 14 above, its coefficient remains significant at a 95% confidence level unless the α_z parameters are restricted to be equal across all seven sectors. If this restriction is imposed, as in the models of Table 2, the coefficient on the filtered SIG variable becomes insignificant.

[26]This bias is not reflected in the tests for serial correlation and heteroscedasticity undertaken above and, thus, such tests may not be good signals of aggregation bias.

Ahmed, Shaghil. "Wage Stickiness and the Non-Neutrality of Money: A Cross-Industry Analysis." *Journal of Monetary Economics* 20 (July 1987): 25–50.

Barro, Robert J. "Unanticipated Money Growth and Unemployment in the United States." *American Economic Review* 67 (March 1977): 101–15.

——. "Unanticipated Money, Output and the Price Level in the United States." *Journal of Political Economy* 86 (August 1978): 549–80.

——. "Unanticipated Money Growth and Economic Activity in the United States." In *Money Expectations and Business Cycles*, edited by R. J. Barro. New York: Academic Press, 1981.

Barro, Robert J., and Mark Rush. "Unanticipated Money and Economic Activity." In *Rational Expectations and Economic Policy*, edited by S. Fischer. Chicago: University of Chicago Press, 1980.

Breusch, Trevor S., and Adrian R. Pagan. "A Simple Test for Heteroscedasticity and Random Coefficient Variation." *Econometrica* 47 (1979): 1287–94.

Chan, Louis K. C. "Unanticipated Monetary Policy and Real Economic Activity: Some Cross-Regime Evidence." *Journal of Monetary Economics* 22 (November 1988): 439–59.

Davidson, Russell, and James G. Mackinnon. *Estimation and Inference in Econometrics.* New York: Oxford University Press, 1993.

Dickey, David A., William R. Bell, and Robert B. Miller. "Unit Roots in Time Series Models: Tests and Implications." *American Statistician* 40 (February 1986): 12–26.

Efron, Bradley, and R. Tibshirani. "Bootstrap Methods for Standard Errors, Confidence Intervals, and Other Measures of Statistical Accuracy." *Statistical Science* 1 (1986): 54–77.

Engle, Robert F. "Autoregressive Conditional Heteroscedasticity with Estimates of the Variance of United Kingdom Inflation." *Econometrica* 50 (1982): 987–1007.

Fair, Ray. "An Analysis of the Accuracy of Four Macroeconometric Models." *Journal of Political Economy* 87 (August 1979): 701–18.

Freedman, David A., and Stephen C. Peters. "Bootstrapping a Regression Equation: Some Empirical Results." *Journal of the American Statistical Association* 79 (1984a): 97–106.

——. "Bootstrapping an Econometric Model: Some Empirical Results." *Journal of Business and Economic Statistics* 2 (1984b): 150–8.

Fuller, Wayne A. *Introduction to Statistical Time Series.* New York: John Wiley and Sons, 1976.

Gauger, Jean A., and Walter Enders. "Money Neutrality at Aggregate and Sectoral Levels." *Southern Economic Journal* 55 (January 1989): 771–78.

Godfrey, L. G. "Testing For Higher Order Serial Correlation in Regression Equations When The Regressors Include Lagged Dependent Variables." *Econometrica* 46 (November 1978): 1303–1310.

Granger, Clive W. J., and Paul Newbold. "Spurious Regressions in Econometrics." *Journal of Econometrics* 2 (1974): 111–20.

Gray, Jo Anna, Magda Kandil, and David E. Spencer. "Some Evidence on the Role of Contractual Wage Rigidities in Determining Output and Employment." May 1987. Mimeo.

Hoffman, Dennis L., and Don E. Schlagenhauf. "An Econometric Investigation of the Monetary Neutrality and Rationality Propositions from an International Perspective." *Review of Economics and Statistics* 64 (1982): 562–71.

Hoffman, Dennis L., Stuart A. Low, and Don E. Schlagenhauf. "Tests of Rationality, Neutrality and Market Efficiency." *Journal of Monetary Economics* 14 (1984): 339–63.

Kretzmer, Peter E. "The Cross-Industry Effects of Unanticipated Money in an Equilibrium Business Cycle Model." *Journal of Monetary Economics* 23 (1989): 275–96.

Leiderman, Leonardo. "Macroeconometric Testing of the Rational Expectations and Structural Neutrality Hypotheses for the United States." *Journal of Monetary Economics* 6 (1980): 69–82.

Lilien, David M. "Sectoral Shifts and Cyclical Unemployment." *Journal of Political Economy* 90 (1982): 777–93.

Lucas, Robert E. "Some International Evidence on Output–Inflation Trade-offs." *American Economic Review* 63 (June 1973): 326–34.

Mishkin, Frederic. "Does Anticipated Aggregate Demand Policy Matter? Further Economic Results." *American Economic Review* 72 (1982a): 788–802.

———. "Does Anticipated Monetary Policy Matter? An Economic Investigation." *Journal of Political Economy* 90 (1982b): 22–51.

———. *A Rational Expectations Approach to Macroeconometrics.* Chicago: University of Chicago Press, 1983.

Neftci, Salih, and Thomas J. Sargent. "A Little Bit of Evidence on the Natural Rate Hypothesis from the U. S." *Journal of Monetary Economics* 4 (1978): 315–19.

Pagan, Adrian. "Econometric Issues in the Analysis of Regressions with Generated Regressors." *International Economic Review* 25 (February 1984): 221–47.

———. "Two Stage and Related Estimators and Their Applications." *Review of Economic Studies* 53 (August 1986): 517–38.

Pesaran, M. Hashem. "A Critique of the Proposed Tests of the Natural Rate–Rational Expectations Hypothesis." *Economic Journal* 92 (September 1982): 254–81.

Rasche, Robert H. "M1—Velocity and Money—Demand Functions: Do Stable Relationships Exist?" In *Empirical Studies of Velocity, Real Exchange Rates, Unemployment and Productivity*, edited by K. Brunner and A. H. Meltzer, 9–88. Carnegie-Rochester Conference Series on Public Policy 27, Autumn 1987.

Rush, Mark. "Unexpected Money and Employment, 1920 to 1983." *Journal of Money, Credit and Banking* 18 (August 1986): 259–74.

Sargent, Thomas J. "A Classical Macroeconometric Model for the United States." *Journal of Political Economy* 84 (April 1976): 207–37.

Theil, Henri. *Principles of Econometrics*. New York: John Wiley and Sons, 1971.

Appendix

Data Sources

Data from the following ten sectors is used: construction; mining; durable manufacturing; non-durable manufacturing; wholesale trade; retail trade; finance, insurance and real estate; services; transportation, communication, electric, gas and sanitary services; and agriculture. In order to reduce the number of parameters to be estimated, the price and output series used were derived after combining durable with non-durable manufacturing, wholesale trade with retail trade, and finance, insurance and real estate with services.

The sectoral price indices and constant dollar output series came from industry level current and constant dollar GNP series. The sources for these are: 1947–1975, *The National Income and Product Accounts of the United States, 1929–1975*, Statistical Tables, U.S. Department of Commerce, Bureau of Economic Analysis, pp. 227–28; 1976, *National Income and Product Accounts, 1976–79*, Survey of Current Business Special Supplement, July 1981, Bureau of Economic Analysis, U.S. Department of Commerce, p. 49; 1977–1985, *Survey of Current Business*, January 1991, Bureau of Economic Analysis, U.S. Department of Commerce, pp. 33–4; 1986–1989, *Survey of Current Business*, April 1991, Bureau of Economic Analysis, U.S. Department of Commerce, p. 27. To accommodate differences in the pre- and post-1977 data, each series in the pre-1977 data was adjusted by a multiple such that the mean of the available overlapping data (1977–1986) was the same for the two sets of data.

Employment by sector, required for the creation of the SIG variable, came from the following sources: 1947–1960, *Business Statistics*, 1971, U.S. Department of Commerce, pp. 67, 69–71; 1961–1984, *Business Statistics, 1984*, U.S. Department of Commerce, pp. 43, 46–48; 1985–1986, *Supplement to Employment and Earnings*, U.S. Department of Labor, Bureau of Labor Statistics, August 1988, for all sectors except agriculture which came from *Employment and Earnings*, Bureau of Labor Statistics, U.S. Department of Labor, February 1989, p. 9; 1987–1989, *Employment and Earnings*, Bureau of Labor Statistics, U.S. Department of Labor, February 1992, p. 47.

The energy price index used in the price equation is the fuel oil and coal consumer price index which came from the following sources: 1947–1960, *Business Statistics*, 1971, Bureau of Economic Analysis, U.S. Department of Commerce, 1972, p. 41; 1961–1984, *Business Statistics, 1984*, Bureau of Economic Analysis, U.S. Department of Commerce, 1985, p. 25; 1985–1986, *Survey of Current Business*, November 1987, Bureau of Economic Analysis, U.S. Department of Commerce, p. 5–6; 1987–1989, *Survey of Current Business*, Bureau of Economic Analysis, U.S. Department of Commerce, various issues.

The monetary aggregate used in the price forecast equation is the annual average of monthly observations on M1. The sources for this are: 1947–1958, *Business Statistics, 1971*, Bureau of Economic Analysis, U.S. Department of Commerce, 1972, adjusted using the method employed in Rasche (1987) to make them consistent with later data; 1959–1986, CANSIM data base, series B54324, Statistics Canada; 1987–1989, *Survey of Current Business*, Bureau of Economic Analysis, U.S. Department of Commerce, various issues.