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Sticky Prices: An Empirical Assessment of Alternative Models

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Abstract

This paper presents a model of staggered price setting that allows for a flexible distribution of the durations of the prices underlying aggregate price behavior, and estimates it with U.S. data. When tested against an unrestricted version of this model, standard models of sticky prices are rejected. In contrast, a stylized model that assumes a trimodal distribution of price durations—with clusters on the first, fourth, and eighth quarter after prices are set—easily passes the same test. In addition, this model is able to replicate the dynamic behavior of inflation and output found in the data.

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I. INTRODUCTION

Since the early 1990s, a sizable amount of research in macroeconomics has focused on building quantitative dynamic-general-equilibrium models that include sticky nominal prices.[2] According to the literature reporting this research, a key factor behind this development has been the attempt to overcome the difficulties that real business cycle models with flexible prices have encountered in explaining the interaction between real and nominal variables found in the data. The increasing popularity of these new breed of models also has been bolstered by the desire to compare the consequences of alternative monetary policy rules on the basis of a framework that is both theoretically rigorous and empirically relevant.[3]

A major difficulty for research in this area is that it is unclear which among a number of approaches to modeling sticky prices is more appropriate for empirical purposes. Following an analysis that Taylor (1980) originally proposed for wages, one popular approximation to the subject is to assume that price decisions are staggered and that the duration of the individual prices resulting from those decisions is uniform and equal to or less than four quarters. Another standard approach relies on the use of the Calvo-Rotemberg model, which in Calvo's (1983) version assumes that price decisions are staggered and that the duration of each price is a random variable distributed according to an exponential function of the interval since the price was set for the last time, while in Rotemberg's (1982) version assumes that individual price setters face a quadratic cost of adjusting prices.[4] Typical quantifications of the Calvo-Rotemberg model imply an average duration of individual prices between three and twelve quarters. A third popular approximation to the subject is to assume that individual prices are revised every period, without any staggering of price decisions, but with nominal

[2]Nelson (1998a) estimates that, since 1990, there have appeared more than 50 papers adding nominal rigidities to otherwise standard real business cycle models. Most of this literature includes quantitative simulations or estimations.

[3]Additional factors may be the improved theoretical understanding of the sources for nominal price rigidities, the growing formal evidence that many prices are not modified during weeks, months, or years, and the continued demand from policymakers for practical models that take these rigidities into account. For a general survey of research on explaining, documenting, and modeling sticky prices, see Taylor (1998).

[4]Despite the different microfoundations of the Calvo (1983) and Rotemberg (1982) models, their aggregate implications are similar. See Rotemberg (1987) and Roberts (1995).

prices being set in advance. Papers following this approach typically specify the basic period of analysis to be equal to one quarter, implying an average duration of individual prices substantially shorter than in the above approaches.[5]

To further complicate matters, a number of recent papers have questioned the ability of the above mentioned models to replicate the dynamic behavior of inflation and output observed in the United States (e.g., Chari, Kehoe, and McGrattan (1996); Fuhrer and Moore (1995), Nelson (1998b)). The key concern of this research is that those models would be unable to generate enough persistence in inflation and output. Specifically, while the evidence indicates that aggregate shocks to those variables typically have had effects that have persisted well beyond the first year following the shock, the simulations and estimates provided by this recent research have suggested that the above mentioned models may not be able to generate movements in inflation and output that persist for more than one year.[6]

This paper assesses the empirical validity of alternative models of sticky prices for the United States. For this purpose, it presents a model of staggered price setting that allows for a flexible distribution of the durations of the individual prices underlying the behavior of the aggregate price, estimates it on the basis of aggregate data for that country, and uses the results to test simple competing models. The paper shows that, when tested against an unrestricted version of the model, the standard approaches used to model sticky prices in the recent literature on quantitative-dynamic-general-equilibrium models are rejected by the data. In contrast, it finds that a stylized model that allows for a trimodal distribution of price durations with clusters on the first, fourth, and eighth quarter after prices are set, easily passes the same test. In addition, the paper shows that this model is able to replicate the main aspects

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[5]For a more detailed description of these and other models of nominal price rigidities used in the economic literature, see Taylor (1998), Nelson (1998a, 1998b), and the references therein.

[6]Chari, Kehoe, and McGrattan (1996) have argued that, in a plausibly parameterized dynamic-general-equilibrium model, a model *à la* Taylor with prices lasting for four quarters fails to generate enough persistence in the movements of output following monetary shocks (their view, however, has been challenged by Erceg, 1997, Jeanne, 1998, and Kiley, 1997). Fuhrer and Moore (1995) have argued that a staggered price setting model with prices lasting up to four quarters cannot reproduce the persistence of inflation and the correlations between output and inflation observed in the data, and go on to propose an alternative model that would allow for more interesting dynamics at the cost of weaker microfoundations (for a discussion on the latter, see Taylor, 1998). More recently, Nelson (1998b) has questioned the ability of standard quantifications of the Taylor and the Calvo-Rotemberg models to replicate the lags in the response of inflation to monetary growth, and has reiterated the concern that a model *à la* Taylor with prices that last for four quarters does not impart enough inflation persistence. Nelson (1998b) also has raised similar objections to models that assume that prices are fixed only for one quarter.

of the dynamic behavior of inflation and output observed in the data, including the degree of persistence of the movements in these variables.

The remainder of this paper is organized as follows. Next section presents the basic model used for the analysis, and shows that the distribution of price durations underlying aggregate price behavior can be estimated using data on inflation and other aggregated variables. Sections III estimates an unrestricted version of this model with quarterly data for the United States. Section IV uses the results of the previous section to test competing parsimonious models of sticky prices for that country. Section V examines the ability of the preferred model to replicate the dynamic behavior of inflation and output found in the data. Section VI offers concluding remarks.

II. A MODEL OF STAGGERED PRICE SETTING [7]

The duration of an individual price is defined as the number of periods during which that price is fixed. This section presents a model of staggered price setting that allows for any distribution of price durations, and shows that this distribution can be inferred on the basis of standard data on inflation and other macroeconomic variables.[8]

The model contains three basic equations. The first one writes the inflation rate, $\hat{p}_t$, as a weighted average of partial inflation indicators that measure the aggregate rate of change of the prices with similar durations:

$$\hat{p}_t = \sum_{s=1}^{n} \lambda_s \hat{p}_t^s, \tag{1}$$

where $\hat{p}_t^s$ is the aggregate rate of change of the prices with duration s=1,2...n, λ_s is the weight of those prices in the aggregate price index (so that $\lambda_s \geq 0$ and the sum of λ_s over s is unity), and the superscript t represents the period for which the associated variable is measured.

The second equation expresses $\hat{p}_t^s$ as a function of the average rate of change of the prices with duration s. For this purpose, it is assumed that the decisions about prices with the same duration are uniformly staggered, so that the fraction of those prices revised in any given

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[7]This section builds on Jadresic (1992), who examines the type of wage contracts underlying the behavior of the aggregate wage in Chile.

[8]Benabou and Bismut (1988), Taylor (1993), and Fuhrer and Moore (1995) also propose and estimate models for the aggregate wage or price that allow for a flexible distribution of the underlying wage or price durations. The models and/or the estimates they present, however, impose ad hoc constraints in the shape of that distribution.

period is 1/s.[9] Also, to simplify notation, it is assumed that the individual prices set at period t with duration s are all identical. Representing the log of these prices by x_t^s, it follows that

$$\hat{p}_t^s = \frac{1}{s}\,(x_t^s - x_{t-s}^s) = \frac{1}{s}\,(1 - L^s)\,x_t^s. \tag{2}$$

where L is the standard lag operator.

The third equation specifies x_t^s using the following approximation to optimal price setting in a multiperiod framework:

$$x_t^s = E_{t-1}\left[\sum_{i=0}^{s-1}\left(\frac{P_{t+i} + \gamma_s Y_{t+i}}{s}\right)\right] + a_t^s, \tag{3}$$

where P_t is the log of the aggregate price level, Y_t the log of aggregate output, E_{t-1} the standard expectations operator conditional on information about variables realized up to t-1, γ_s a positive parameter, and a_t^s is an exogenous stochastic variable that captures the effects of microeconomic factors on the price set at time t. This equation can be interpreted as a log-linear approximation to the optimal pricing rule of a monopolistically competitive producer that sets its price for the period [t,t+s) on the basis of information of variables realized up to t-1, under the simplifying assumption of no time discounting. Intuitively, it states that, apart from the microeconomic factors that determine a_t^s, price setters care about the average aggregate price and output expected to prevail until the next price revision.[10]

Replacing (3) in (2) and (1) implies

$$\hat{p}_t = \sum_{s=1}^{n} \lambda_s \frac{1}{s}\,(1 - L^s)\,E_{t-1}\left[\sum_{i=0}^{s-1}\left(\frac{P_{t+i} + \gamma_s Y_{t+i}}{s}\right)\right] + \epsilon_t, \tag{4}$$

where ϵ_t is the sum of $\lambda_s(1/s)(1\text{-}L^s)a_t^s$ over s=1..n.

The right hand side of equation (4a) can be rewritten so that inflation is expressed as a linear function of a set of lagged inflation rates, expected inflation rates, and expected levels

[9]Jadresic (1992) finds that relaxing this assumption can be important in order to assess the effects of seasonality. The analysis below simplifies by focusing on seasonally adjusted data.

[10]Note that the assumption of no time discounting can be relaxed by introducing time discounting factors in the right hand side of equation (3). The empirical validity of this assumption is examined in Section III.

of aggregate output. Specifically, inserting $P_{t+i}=P_{t+i-1}+\hat{p}_{t+i}$ for all $i\geq 0$, using the fact that $(1-L^s)P_{t-1}$ is equal to the sum of $\hat{p}_{t-i}$ from i=1..s, and rearranging terms, it follows that

$$\hat{p}_t = \sum_{s=1}^{n} \lambda_s \sum_{i=1}^{s} \frac{\hat{p}_{t-i}}{s} + \sum_{s=1}^{n} \lambda_s \frac{1}{s} (1-L^s) \sum_{i=0}^{s-1} \left(\frac{(s-i)\, E_{t-1}\hat{p}_{t+i} + \gamma_s E_{t-1} Y_{t+i}}{s} \right) + \epsilon_t. \tag{4b}$$

Equation (4b) is the basis for the analysis in the remainder of this paper. The intuition is as follows. The first term in the right hand side represents the effect on current inflation of the catch-up price increases designed to reestablish the value of the prices currently being revised to the real level they had when they were set the last time. The second term corresponds to the effect on current inflation of the component of current price revisions that is attributable to changes in the price-setters' expectations about future inflation and aggregate output. As indicated by the term $(1-L^s)$ appearing in this term, the relevant change in expectations is the change occurred since the last time that the prices currently being revised were set. The third term corresponds to the aggregate effect of the microeconomic factors affecting price setting.

Equations (4b) or (4a) can be used for various purposes. For instance, if one is willing to make a-priori assumptions about the distribution of the price durations (λ_s for s=1...n), the partial elasticities of prices with respect to output (γ_s), and the stochastic process determining a_t^s or ϵ_t, this equation provides a full characterization of aggregate price setting, one that can be embedded in dynamic-general-equilibrium models for simulation purposes. Alternatively, and more importantly for this paper, if one has data or a model for expected inflation and aggregate output, and makes an auxiliary assumption about the stochastic process a_t^s or ϵ_t, equation (4b) can be used to estimate econometrically the distribution of price durations, as well as the partial elasticities of prices relative to output.

To illustrate, suppose that aggregate output is determined by real money balances, and that nominal money balances follow a random walk; i.e.,

$$Y_t = M_t - P_t, \tag{5}$$

$$M_t = M_{t-1} + \mu_t, \tag{6}$$

where M_t is the log of money balances, and μ_t is white noise. In addition, suppose that $\gamma_s=1$ for all s. Assuming that inflation and output expectations are model-consistent, it follows that

$$E_{t-1}\left[\sum_{i=0}^{s-1}\left(\frac{P_{t+i}+\gamma_s Y_{t+i}}{s}\right)\right] = M_{t-1} \tag{7}$$

This simple model for inflation and output expectations can be used in equation (4a) to obtain an equation for inflation as a function of past changes in money balances. Denoting money growth by $\hat{m}_t$, the latter also can be written as

$$\begin{aligned}\hat{p}_t &= \sum_{s=1}^{n} \lambda_s \sum_{i=1}^{s} \frac{\hat{m}_{t-i}}{s} + \epsilon_t \\ &= \lambda_1 \hat{m}_{t-1} + \lambda_2\left(\frac{\hat{m}_{t-1}+\hat{m}_{t-2}}{2}\right) + \ldots + \lambda_n\left(\frac{\hat{m}_{t-1}+ .. +\hat{m}_{t-n}}{n}\right) + \epsilon_t .\end{aligned} \tag{8}$$

If ϵ_t can be modeled as white noise, equation (8) implies that the distribution of price durations λ_s can thus be inferred from the coefficients of a standard linear regression of inflation on lagged rates of money growth. For instance, if the coefficient on the first lag of money growth resulting from that regression is equal to one, while all the remainder coefficients are nil, it can be deduced that prices last for only one period. Alternatively, if the coefficients on the first and second lag of money growth are both equal to one half, while the remainder coefficients are nil, it follows that prices last for two periods. Or if the findings are that the coefficient on the first lag of money growth is 0.75, the coefficient on the second lag is 0.25, and the remainder coefficients are nil, then it can be inferred that half of the prices last for one period, while the other half last for two periods. In general, any underlying distribution of price durations can inferred by proper comparison of the regression coefficients with the ones implied by equation (8) for alternative distributions of price durations.

Of course, this example is based on a model for output and inflation expectations that is too simplistic for actual empirical work. Nonetheless, it should suffice to illustrate the general principle that the dynamic behavior of the aggregate data on inflation and other macroeconomic variables provides information about the distribution of price durations underlying that data. The intuition behind this principle is that the adjustment of the prices that are revised frequently depends only on recent news about the fundamental factors that determine prices, while the adjustment of prices that are revised infrequently also depends on older news about those factors.

Before applying this principle to actual data, note that equation (4b) also provides a benchmark for assessing the empirical validity of competing models of sticky prices. In particular, the approach that assumes that prices are revised every four quarters corresponds to the special case λ_4=1 (and λ_s=0 for all s≠4). Also, the one quarter pricing approach corresponds to the case λ_1=1 (and λ_s=0 for all s≠1). Thus these and other simple approaches to modeling sticky prices can be tested by examining the empirical validity of the restrictions that those approaches impose on the coefficients λ_s.

One approach that cannot be directly associated to a particular case of equation (4b) corresponds to the Calvo-Rotemberg model. It can be shown, however, that, if individual prices are set on the basis of information about variables realized up to t-1, the Calvo-Rotemberg model implies a price adjustment equation of the type

$$\hat{p}_t = E_{t-1}\hat{p}_{t+1} + \gamma\, E_{t-1} Y_t + b_t, \tag{9}$$

where b_t is an exogenous stochastic variable analogous to a^s_t.[11] The empirical usefulness of the Calvo-Rotemberg approach relative to the staggered price setting model presented above can be examined by expressing the inflation rate as a weighted average of the right hand sides of equations (4b) and (9) and testing the statistical significance of the weights associated to those two components.

III. Estimation and Tests of the Basic model

This section describes the data and procedure used to estimate equation (4b) on the basis of aggregate data from the United States, and presents and tests the implied results.

Following Fuhrer and Moore (1995), the empirical analysis is conducted with quarterly data for the nonfarm business sector, with $\hat{p}_t$ and Y_t being measured respectively by the change in the log of the implicit deflator (multiplied by four to express quarterly inflation in annual terms) and the level of the log of real output in that sector. The series are seasonally adjusted, and are taken from the DRI macroeconomics database, formerly the Citibank database.

The base estimates are based on a data sample for the period 1979:4-1998:2. This is the most natural period for the estimations in this paper, since previous studies have found that there was a significant difference in the way monetary policy was conducted pre and post October 1979; i.e., the VAR equations used to estimate inflation and output expectations cannot be assumed to have been the same in earlier periods.[12] For the sake of comparison,

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[11] See Roberts (1995) for the derivation under the assumption that individual prices are set on the basis of information available about variables realized up to period t; the extension to the case considered here is trivial.

[12] On the evidence of the changes in the Fed's monetary policy reaction function, see for instance Clarida, Gali and Gertler (1998), and Judd and Rudebusch (1998).

however, the analysis also was conducted for an extended sample stretching back to 1975:2.[13] Note that, when using either of these samples, estimating equation (4b) under the assumption that the price durations may be up to eight quarters implies loosing eight initial observations.

To model the inflation and output expectations appearing in the right hand side of equation (4b), an auxiliary vector autoregression (VAR) model was estimated including inflation, output, and the short-term nominal interest rate. As argued by Fuhrer and Moore (1995), and documented by Bernanke and Blinder (1992), the latter variable is an important predictor of future output, and is thus relevant for estimating output expectations. The particular measure used for the short-term nominal interest rate was the Federal Funds rate. The expected values of inflation and output were obtained by forecasting those variables with the VAR model at the different relevant horizons on the basis of the variables realized in previous periods. The VAR model used to obtain the base estimates included eight lags of each of the variables. The results implied by considering other lag lengths, and by replacing in the VAR model the short-term nominal interest rate by the monetary base, are reported below.

For estimation purposes, the stochastic variable ϵ_t is assumed to be a constant plus white noise, which can be justified from the more basic premise that the microeconomic component of a price of duration s is determined according to a random walk of order s, with drift (i.e., $a^s_t = a^s_{t-s}+\alpha_s$+white noise, where α_s is a parameter). This specification for ϵ_t has the advantage that it puts the burden of the statistical explanation of inflation on the variables explicitly taken into account in the analysis, rather than on a more complicated and harder to interpret exogenous stochastic process.[14] The base estimates also assume that the partial elasticity of prices with respect to output is independent of the price durations ($\gamma_s=\gamma$ for all s), and that the maximum relevant price duration is eight quarters (n=8). The statistical validity of these auxiliary hypothesis about ϵ_t, γ_s, and n is examined below.

Table 1 reports the results of estimating equation (4b) for the base specification and sample described in the above paragraphs. All the regressions reported in this table and in the remainder of the paper were estimated using nonlinear least squares. The first two regressions are included only to provide background information: regression (A) imposes no constraints on the estimates for the weights λ_s, while regression (B) only restricts them to sum unity.

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[13]Extending the sample period even further into the past would raise problems of interpretation and technique because during Nixon's government there were widespread price controls that distorted significantly aggregate price behavior until early 1975. Taking 1975:2 as the first quarter with fully undistorted prices is suggested both by the evidence in Frye and Gordon (1980), Gordon (1981), and Meyer (1982), and by the hypothesis that the lagged effects of Nixon's price controls on the aggregate price lasted for up to eight quarters.

[14] On this issue, see Fuhrer's (1997) and McCallum's (1997) comments on Rotemberg and Woodford (1997)

Table 1. Estimates of the Basic Model
(Standard errors in parenthesis)

Regression	Constant	λ_1	λ_2	λ_3	λ_4	λ_5	λ_6	λ_7	λ_8	γ	R^2	$\bar{R}^2$	SER $\times 10^{-3}$	SSR $\times 10^{-3}$	DW	Average Duration
(A)	0.0037	0.20	0.03	-0.04	0.71	-0.37	-0.46	0.41	0.36	0.13	0.71	0.67	7.6	3.32	2.05	4.1
	(0.0035)	(0.13)	(0.26)	(0.34)	(0.41)	(0.45)	(0.51)	(0.56)	(0.39)	(0.23)						(0.9)
(B)	-.0026	0.23	0.00	0.02	0.67	-0.32	-0.45	0.42	0.43	0.31	0.69	0.65	7.8	3.55	2.01	5.0
	(0.0017)	(0.14)	(0.27)	(0.35)	(0.42)	(0.47)	(0.54)	(0.59)	(0.41)	(0.19)						(0.8)
(C)	-0.0027	0.27	-	0.04	0.29	-	-	-	0.40	0.32	0.68	0.66	7.7	3.71	2.07	4.8
	(0.0017)	(0.12)		(0.27)	(0.30)				(0.16)	(0.19)						(0.8)
(D)	-0.0027	0.28	-		0.32		-	-	0.40	0.32	0.68	0.66	7 .7	3.71	2.07	4.8
	(0.0017)	(0.11)			(0.20)				(0.16)	(0.17)						(0.7)

Notes: Based on 67 quarterly observations for 1981:4-1998:2. The estimates correspond to equation (4b) in the text, with γ_s=γ for all s. Regression (A) allows for the weights λ_s to be negative and for their sum to be different from unity. Regression (B) allows for the weights λ_s to be negative but imposes the restriction that their sum is unity. Regression (C) imposes the model's restrictions that $\lambda_s \geq 0$ and that the sum of the weights λ_s is unity. Regression (D) assumes, in addition, that λ_3 is nil.

Since several of the weights implied by these regressions are negative, those regressions do not provide valid estimates of the model. For this reason, the core estimates of the model are those associated to regression (C), which constrains the estimated weights to be nonnegative and their sum to be equal to unity.[15]

Regression (C) shows the key result of this paper: minimizing the sum of squared residuals implies assigning practically all of the weight of the explanation of the conduct of the aggregate price to the behavior of individual prices with durations of one, four, and eight quarters. In contrast, the same criteria implies dropping from that explanation the behavior of prices with durations equal to two, five, six, and seven quarters. It also implies giving a very small weight to the effect of prices with durations of three quarters, which for parsimony can be set equal to zero to obtain an equation like regression (D).

Before further examining the implications of this result for the appropriate characterization of the distribution of price durations, it is useful to examine these regressions in more detail. At this respect, several remarks are worth making.

First, the remainder estimates implied by these regressions seem reasonable from an economic viewpoint. In particular, the estimated partial elasticity of prices with respect to output of about 0.3 has the correct sign, and an order of magnitude that is economically significant, with a t-statistic above unity. To compare, Fuhrer and Moore's (1995) estimate of a similar elasticity is of 3.2×10^{-7} with a t-statistic of 0.06. In addition, the regressions imply an average price duration between four and five quarters. This estimate is highly agreeable with Taylor's (1998) assessment of the microeconomic evidence on wage and price setting, which in his view suggests an average price duration in the United States of about one year.[16]

Second, the regressions seem satisfactory from a statistical viewpoint. To obtain a benchmark to evaluate the goodness of the fit, a series of autoregressive models of inflation with one to twelve lags were estimated, as well as a series of multivariate regressions of

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[15]These constraints were imposed by estimating equation (4b) with all but one of the weights λ_s replaced by the square root of their squared values, substituting the remainder weight with one minus the sum of the other weights, and checking that the implied estimate for the substituted weight was nonnegative. Since estimation by least squares using the Gauss-Newton algorithm in Rats provided slightly different results depending on which was the specific weight eliminated from the original equation and on the initial values provided for the estimation, the coefficients of Regression (C) were estimated using instead Rats' simplex search procedure, which provided robust results. Since this procedure does not provide standard errors for the estimated coefficients, the latter were estimated using the standard Gauss-Newton method after dropping from the original equation the weights estimated to be zero when using the simplex method.

[16] The average price duration is defined as $\sum\lambda_s s$, with the sum specified over s=1..8.

current inflation on a constant and lagged inflation rates, output levels, and interest rates, including between one and twelve lags of each variable. For all these models, the adjusted R^2 was found to be of the order of 0.6, and in all cases below 0.63. In comparison, the adjusted R^2 in the regressions in Table 1 are not only comparable to those magnitudes, but actually larger, around 0.66.

Third, the regression results imply that the model passes a number of relevant statistical tests without difficulty. This is shown in Table 2, which reports the results of F tests for the hypotheses that the weights λ_s are nonnegative and their sum is unity, that the partial elasticity of prices with respect to output is independent of the duration of prices, and that price durations above eight quarters are irrelevant. It also shows the results of applying the Breusch-Godfrey test to the hypothesis that the residuals present no serial autocorrelation (see Maddala, 1988). None of the null hypotheses of the model is rejected at standard levels of significance.

Finally, estimating equation (4b) for the alternative sample and specifications considered in Table 3 either does not change the essence of the results or harms the statistical performance of the model. In particular, extending the estimation period to 1977:2-1998:2, modeling inflation and output expectations on the basis of a VAR model with twelve instead of eight lags, modifying equation (4b) in order to include a plausible intertemporal rate of discounting, or modeling inflation and output expectations with a VAR model that includes the monetary base instead of the short-term nominal interest rate, all have only marginal effects on the estimated parameters, and in the former two cases introduces negative serial autocorrelation of the residuals. Modeling inflation and output expectations on the basis of a VAR model with four lags, in turn, considerably worsens the fit of the model, and introduces positive serial autocorrelation in the residuals.[17]

IV. Estimation and Tests of some Parsimonious Models of Sticky Prices

The results of the preceding section indicate that a satisfactory model of sticky prices for the United States is one with a trimodal distribution for the durations of the prices underlying aggregate price behavior, with clusters in the first, fourth, and eighth quarter after prices are set. As shown by the standard errors reported in Tables 1 and 3, however, the specific weights that should be associated to these price durations are imprecisely estimated, raising the question whether a simpler pricing model may provide a reasonable statistical approximation to the data. This section estimates and tests the empirical validity of some simple models of sticky prices, using as standard of reference the basic model estimated above.

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[17]When using the Breusch-Godfrey test, the hypotheses of no first order autocorrelation of the residuals of regressions (E), (G), and (H) are respectively rejected at the 5, 10, and 10 percent significance level.

Table 2. Tests of the Basic Model

Null Hypothesis	Alternative Hypothesis	F test	F Critical Value 5 percent	F Critical Value 1 percent	χ^2 test	χ^2 Critical Value 5 percent	χ^2 Critical Value 1 percent
Weights λ_s are nonnegative and their sum is unity	Some λ_s are negative and/or the sum of the λ_s' is different from unity	1.3	2.4	3.3	-	-	-
Residuals are serially uncorrelated	Residuals are generated by an AR(1) or MA(1) process	-	-	-	0.3	3.8	6.6
Residuals are serially uncorrelated	Residuals are generated by an AR(4) or MA(4) process	-	-	-	0.8	9.5	13.3
Partial elasticity of prices with respect to output is independent of duration of prices (γ_s=γ)	Partial elasticity of prices with respect to output is different for each duration of prices	0.2	2.8	4.1	-	-	-
Price durations above eight quarters are irrelevant (λ_s=0 for s>8 quarters)	Price durations between nine and twelve quarters also should be considered	0.0	2.5	3.7	-	-	-

Note: Based on tests on regression (C) of Table 1.

Table 3. Alternative Estimates of the Basic Model
(Standard errors in parenthesis)

Regression	Constant	λ_1	λ_2	λ_3	λ_4	λ_5	λ_6	λ_7	λ_8	γ	R^2	$\bar{R}^2$	SER $\times 10^{-3}$	SSR $\times 10^{-3}$	DW	Average Duration
(E)	-0.0026	0.31	-	0.01	0.30	-	-	-	0.39	0.37	0.89	0.88	8.7	6.07	2.29	4.7
	(0.0015)	(0.13)		(0.22)	(0.25)				(0.15)	(0.17)						(0.8)
(F)	-0.0013	0.26	-	0.02	0.29	-	-	-	0.42	0.29	0.69	0.66	7.6	3.61	2.07	4.9
	(0.0016)	(0.12)		(0.28)	(0.31)				(0.16)	(0.19)						(0.8)
(G)	-0.0055	0.08	-	0.08	0.31	-	-	-	0.54	0.68	0.61	0.58	8.5	4.48	1.76	5.9
	(0.0022)	(0.12)		(0.25)	(0.27)				(0.15)	(0.25)						(0.7)
(H)	-0.0019	0.37	-	0.04	0.30	-	-	-	0.29	0.22	0.73	0.71	7.1	3.14	2.24	4.0
	(0.0014)	(0.12)		(0.27)	(0.29)				(0.14)	(0.15)						(0.7)
(I)	-0.0026	0.29	-	-	0.31	-	-	-	0.40	0.27	0.65	0.63	8.0	4.05	2.05	4.7
	(0.0018)	(0.12)			(0.18)				(0.15)	(0.21)						(0.8)

Notes: Regression (E) is based on data for the period 1977:2-1998:2. Regression (F) incorporates an intertemporal discounting rate of 8 percent in annual terms. Regression (G) uses the expected values of inflation and output implied by a VAR model with four lags. Regression (H) uses the expected values of inflation and output implied by a VAR model with twelve lags. Regression (I) uses the expected values implied by a VAR model incorporating the monetary base instead of the short-term nominal interest rate.

Table 4 contains the results of estimating four parsimonious models of sticky prices for the same period used to obtain the base estimates of the basic model. The first two regressions report the findings corresponding to the use of the standard approaches that assume that prices last, respectively, for four quarters, and for one quarter (which as noted in Section II can be seen as special cases of the base model). The third regression shows the results of estimating the Calvo-Rotemberg model (equation (9)). The last regression reports the estimates of an alternative model suggested by the results of the preceding section. The latter assumes that the distribution of price durations is trimodal and symmetric, in the sense that the weights corresponding to price durations of one, four, and eight quarters are all set equal to one third. For short, below we will refer to this model simply as the trimodal model.[18]

The first three rows of Table 4 indicate that the statistical performance of the standard models of sticky prices is poor compared to that of the basic model. The results for the Calvo-Rotemberg model, and for the approach that assumes that prices last for one quarter, are particularly unsatisfactory, both in terms of their capability to fit the data and the serial properties of the residuals they imply. The approach that assumes that prices last for four quarters delivers the best fit of the data among the standard models, but this is still considerably inferior to the fit conveyed by the basic model. Moreover, this approach also implies considerable positive autocorrelation in the residuals.

The last row of Table 4 shows that the trimodal model of price durations provides instead a highly satisfactory characterization of the data. In contrast with the standard models, the trimodal model is able to fit the data as efficiently as the basic model, even without adjusting for degrees of freedom. Moreover, the serial autocorrelation of its residuals is small in magnitude, and statistically negligible when applying the Breusch-Godfrey test at the standard levels of significance.

Table 5 provides formal tests of the above parsimonious models using as standard of reference the performance of the basic model. For testing most of these models, the table reports the actual and critical values of F tests for the hypothesis that the particular model being examined imposes valid constraints on the basic model. To assess the empirical validity of the Calvo-Rotemberg model, however, the table reports the results of testing the Calvo-Rotemberg and the basic models against a more general model defined as a weighted average of the basic and the Calvo-Rotemberg models.

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[18]Given the imprecision of the estimates for λ_4, one could posit an even simpler model that includes price durations of one and eight quarters only. The analysis in this and the following section focuses on the trimodal model because it involves λ_s's that are closer to the point estimates of the basic model—which makes the trimodal model relatively harder to reject—and because, a priori, a model that allows also for price durations of four quarters seems more plausible.

Table 4. Estimates of Some Parsimonious Models of Sticky Prices
(Standard errors in parenthesis)

Model	Constant	λ_1	λ_4	λ_8	γ	R^2	$\bar{R}^2$	SER $\times 10^{-3}$	SSR $\times 10^{-3}$	DW
Four quarters prices	-0.0028 (0.0017)	-	1	-	0.34 (0.18)	0.61	0.60	8.3	4.52	1.59
One quarter prices	-0.0015 (0.0015)	1	-	-	0.20 (0.13)	0.44	0.43	9.9	6.4	2.91
Calvo-Rotemberg	0.0064 (0.0720)	-	-	-	0.00 (0.01)	0.47	0.47	9.6	6.04	1.72
Trimodal	-0.0024 (0.0015)	1/3	1/3	1/3	0.29 (0.17)	0.68	0.67	7.6	3.73	2.19

Notes: Based on 67 quarterly observations for 1981:4-1998:2. The values for the λ_s's are imposed a priori. The Calvo-Rotemberg model corresponds to equation (9) in the text.

The results in Table 5 indicate that, using customary significance levels, the standard models of sticky prices are rejected by the data. As shown in the first two rows of the table, this conclusion is direct when testing the approaches that assume that prices last, respectively, for four quarters, and for one quarter. In the case of the Calvo-Rotemberg model, its rejection follows from the results in the third and fourth rows, which reveal that the Calvo-Rotemberg model is rejected when tested against the weighted average model, and that the basic model is statistically equivalent to the weighted average model. Because of the latter, the rejection of the Calvo-Rotemberg against the weighted average model can also be interpreted as its rejection against the basic model.[19]

Finally, the last row of Table 5 shows that, in contrast with the results obtained when testing the standard models of sticky prices, the restrictions imposed by the trimodal model of price durations on the basic model are easily accepted by the data.

V. Replicating the Dynamic Behavior of Inflation and Output

As noted in the Introduction, a number of recent papers have questioned the ability of the standard models of sticky prices to replicate the dynamic behavior of inflation and output observed in the data, specially the persistence in the movements of both variables. This section examines whether the trimodal model proposed in the previous section can reproduce that behavior.

A. Inflation and Output Persistence: A Preliminary Assessment

To start the discussion on familiar grounds, it is useful to examine first the consequences of the trimodal model when output is determined by the aggregate demand schedule

$$Y_t = -P_t + v_t, \tag{9}$$

which can be derived from the quantity theory of money and the assumption that the monetary authority targets constant money balances. We shall focus on the effects of a white-noise demand shock μ_t such that

$$v_t = v_{t-1} + \mu_t, \tag{10}$$

[19]The statistical equivalency of the weighted average and basic models is reflected on the fact that the weight for the Calvo-Rotemberg component of the weighted average model was estimated to be -0.02 (with standard error equal to 0.2) when no constraints were imposed on that weight, and to be nil when a nonnegativity constraint was imposed.

Table 5. Tests of Some Parsimonious Models of Sticky Prices

Null Hypothesis True Model is:	Alternative Hypothesis True Model is:	F test	F critical Value	
			5 percent	1 percent
Model with four-quarters prices	Basic model	4.5	2.8	4.1
Model with one-quarter prices	Basic model	13.1	2.8	4.1
Calvo-Rotemberg model	Weighted average of basic model and Calvo-Rotemberg model	7.5	2.4	3.3
Basic model	Weighted average of basic model and Calvo-Rotemberg model	0.0	3.2	5.0
Trimodal model	Basic model	0.1	2.8	4.1

Note: Based on the results for the regressions reported in Table 4 and regression (C) of Table 1.

which can be interpreted as an unexpected once and for all increase in the velocity of money. Note that the demand shock μ_t also could be interpreted as stemming from an unexpected once and for all upward jump in money supply.

Figure 1 shows the response of inflation and output to such a demand shock in the trimodal model estimated in the previous section. The paths for these variables were obtained by solving by simulation equations (9), (10) and (4b) for the estimates and hypothesis of that model, under the assumption of model-consistent expectations. Not surprisingly, the broad direction of the movements in both variables is as in any standard sticky-price model: initially inflation remains unchanged and all the effect of the shock shows up as an expansion of output, while afterwards inflation rises temporarily and output returns to its original equilibrium. The really important result lies on the considerable persistence shown by the movements in these variables following the shock. In particular, inflation and output remain above their trend values for several quarters after the first year following the shock.

Why is the trimodal model able to generate the degree of persistence in inflation and output observed in Figure 1? One obvious reason is that this model includes a non negligible fraction of prices that last for eight quarters, so that it takes two years for completing the revision of all the individual prices underlying the aggregate price. Another key reason relates to the magnitude of the elasticity of prices with respect to output estimated above. As noted among others by Chari, Kehoe, and McGrattan (1996), the degree of output and inflation persistence when price decisions are staggered depends significantly on the size of that elasticity, with a smaller elasticity implying that prices are less responsive to aggregate fluctuations in output and thus that the adjustment process is slower. The elasticity of the order of 0.3 found in the previous section falls in the lower end of the values deemed plausible by Chari, Kehoe and McGrattan's (1996) on the basis of their preferred calibrations of some simple theoretical models.[20]

B. Replicating the Dynamic Behavior of Inflation and Output Found in the Data

While the results shown in Figure 1 are suggestive, a proper empirical assessment of the ability of the trimodal model to replicate the dynamic behavior of inflation and output demands some further analysis. In particular, such an assessment requires a more precise characterization of the behavior of output and inflation found in the data. In addition, it requires a model of output determination empirically more defensible than the model comprised by equations (9) and (10), which was introduced only for illustrative purposes.

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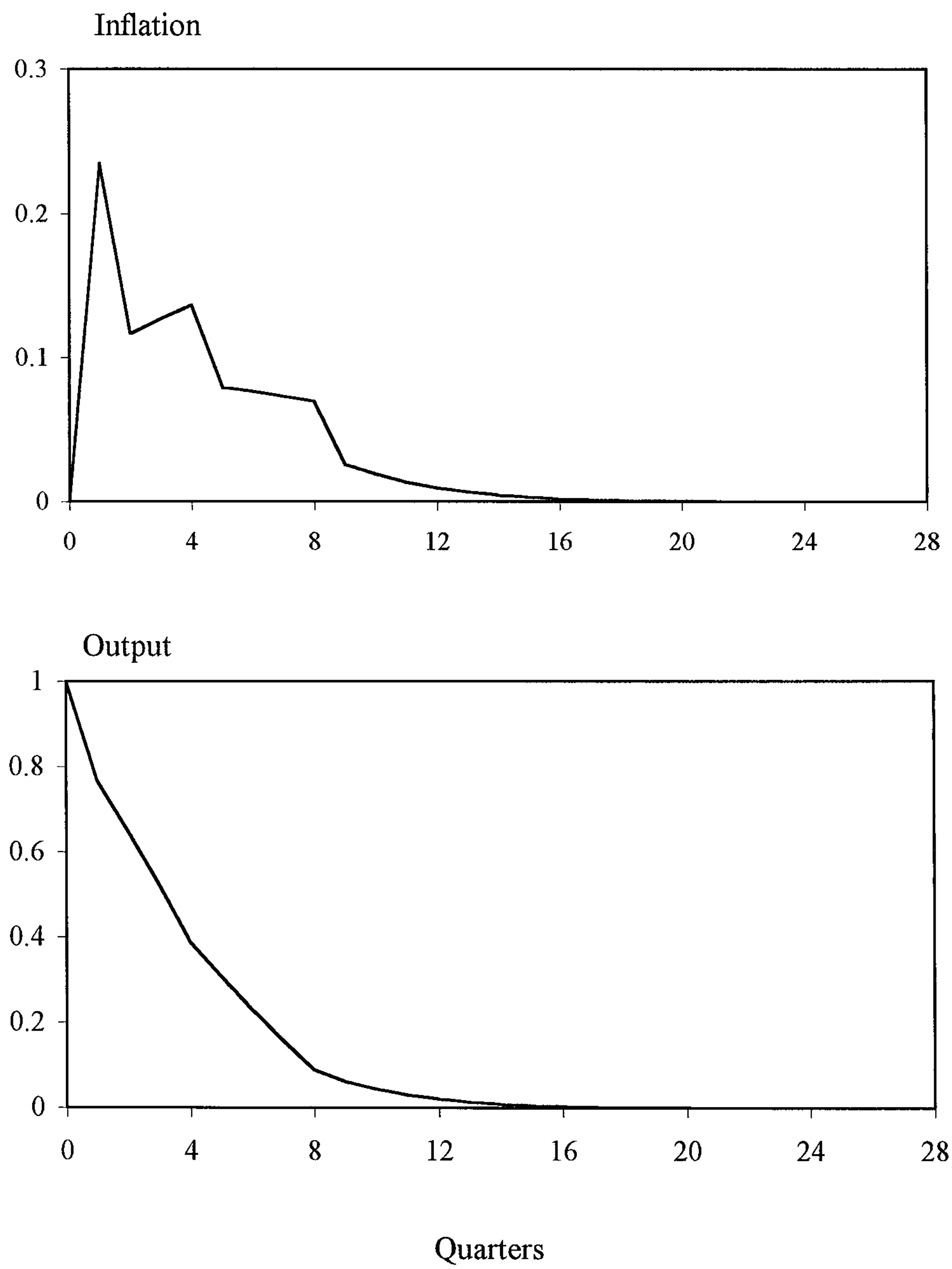
[20]Erceg (1997), Jeanne (1998), and Kiley (1997) provide a number of reasons why γ can be smaller than conjectured by Chary, Kehoe and McGrattan (1996).

Figure 1. Trimodal Model of Price Durations: Effects of a Demand Shock with a Simple Model for Output



To obtain a more satisfactory benchmark for assessing the capacity of the trimodal model to replicate the actual behavior of inflation and output, a bivariate 4-lags VAR model was estimated and used to obtain impulse-response functions for both variables. The impulse-response functions were estimated by orthogonalizing the model's innovations with the Cholesky factorization. Consistent with the sticky-price assumption that prices are set in advance, in this factorization, inflation shocks were ordered first, and output shocks second.[21] The model was estimated for the period 1979:4-1998:2.

Figure 2 shows the impulse-response functions implied by this VAR model. The charts summarize much of the conventional wisdom about the dynamic behavior of output and inflation. Movements in inflation and output tend to persist for significantly more than one year. Also, high inflation tends to lead low output, while high output tends lead high inflation.[22] In the analysis that follows, these impulse-response function are taken to be the stylized facts about the dynamic behavior of inflation and output that the trimodal model should be able to replicate.

As note above, a proper assessment of the ability of the trimodal model to replicate the dynamic behavior of inflation and output also requires considering a more realistic equation for output. For addressing this issue, the price adjustment equation implied by the trimodal model was combined with the output equation implied by the same bivariate VAR model estimated above. While the latter equation cannot be considered as a structural model for output, its use seems sensible for our purposes, since all that is needed is a reasonable summary characterization of the behavior of output in the United States, given the type of monetary policy and behavior of prices observed in that country in the 1980s and 1990s. In addition, using the same output equation in the analysis of the trimodal model and the VAR model has the virtue that it puts the responsibility for the potential discrepancies in the results implied by these models exclusively on the differences on the equations for price adjustment, which are the important equations for the analysis in this paper.

Figure 3 shows the impulse-response functions implied by the trimodal model when combined with the VAR output equation, under the assumption of model-consistent expectations. The ability of the trimodal model to replicate the stylized facts about the dynamic behavior of inflation and output is remarkable. The visual comparison of Figure 2 and Figure 3 indicates that the responses of inflation and output to shocks in both variables is similar in the VAR model and when using the trimodal model of price durations, whether

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[21]Reverting the ordering of the innovations implied very similar results, due to the small correlation found between the residuals of the estimated inflation and output equations.

[22]Note that, under the sticky-price assumption that prices are set in advance, the inflation shock can be interpreted as a supply shock. Moreover, given that in Figure 2 the long-term effect of the output shock on the level of output is virtually nil, it also follows that during the sample period the output shocks can be interpreted as stemming from a pure demand shock.

Figure 2. Impulse-Response Functions for the VAR Model

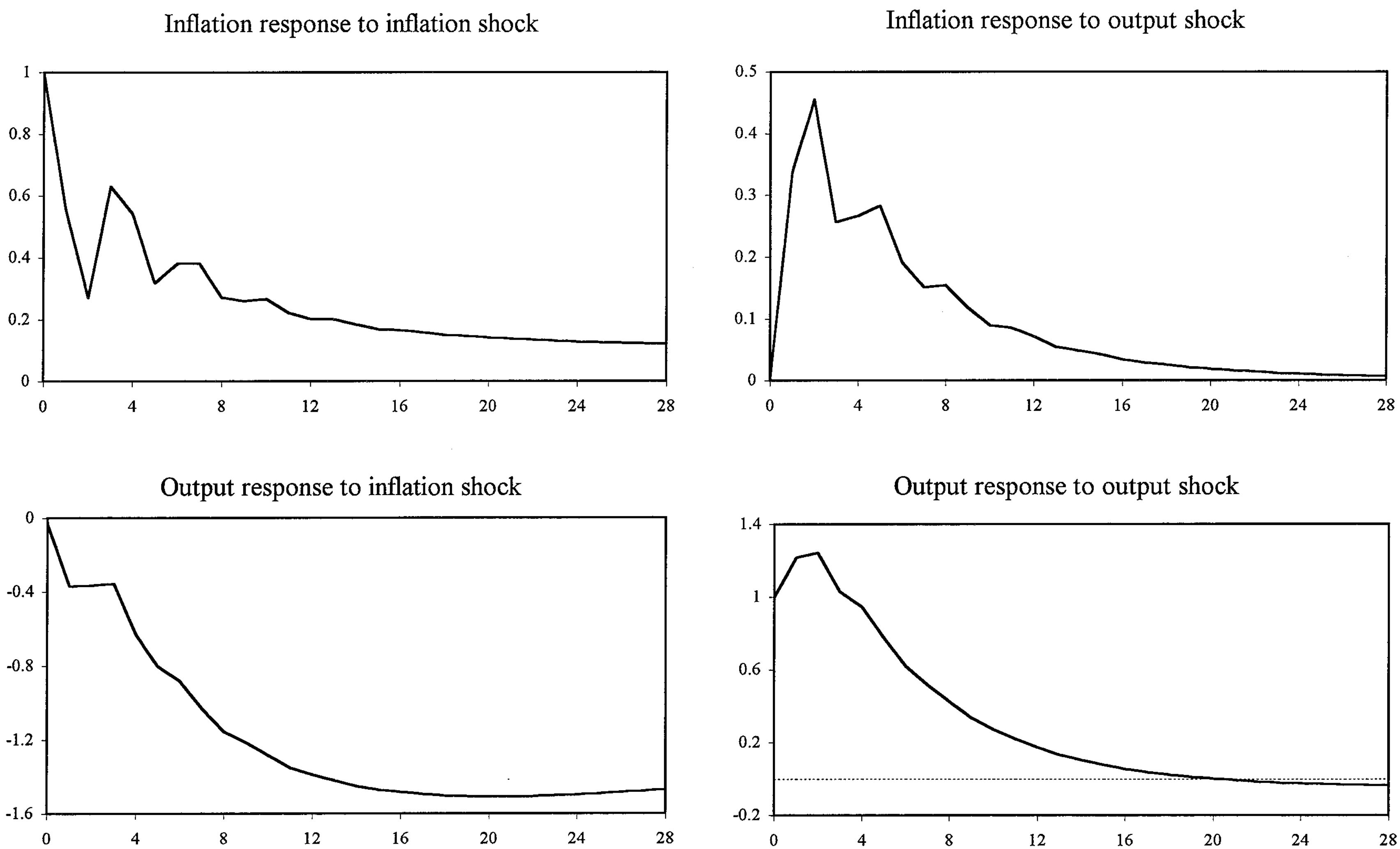
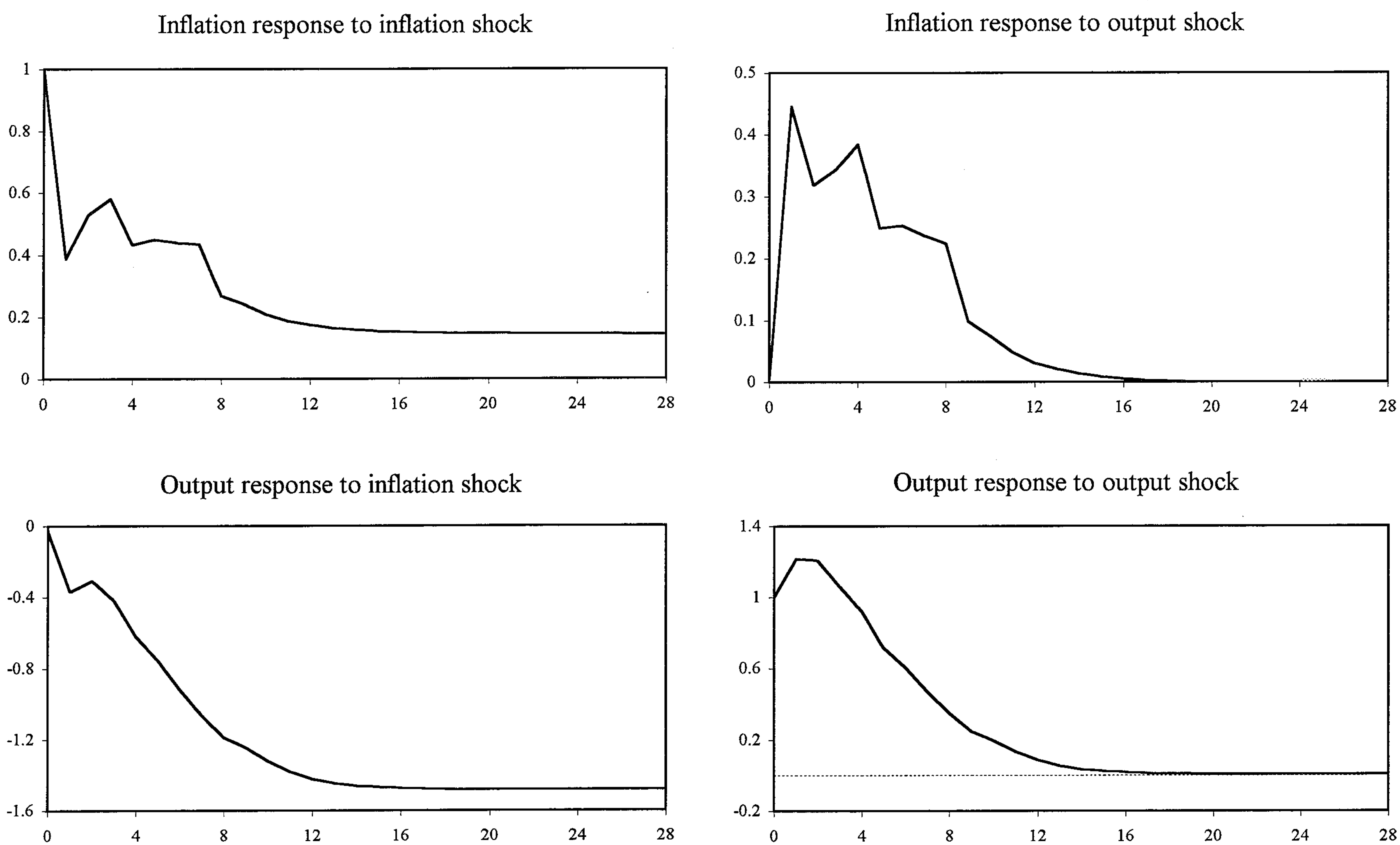


Figure 3. Impulse-Response Functions for the Trimodal Model



measured in terms of the sign, magnitude, or persistence of the effects. Numerically, this similarity is confirmed by the result that the simple correlation between the impulse-response series implied by both models is in all cases above 0.9.

VI. Concluding Remarks

In order to improve the profession's understanding about the actual behavior of real and nominal variables, much of the research in macroeconomics during the 1990s has moved towards the introduction of sticky prices in otherwise standard real business cycle models. It has been unclear, however, which among a number of approaches for modeling sticky prices is more appropriate from an empirical viewpoint. Moreover, some papers have questioned the ability of the standard models of sticky prices to replicate the dynamic behavior of inflation and output in the United States, especially the degree of persistence of the movements in these variables.

This paper presented and estimated a model of staggered price setting that allows for a flexible distribution of price durations, and showed that, unlike standard models of sticky prices, a simple trimodal model of price durations with clusters in the first, fourth and eighth quarter after prices are set provides a satisfactory empirical approximation when using data for the United States. In addition, the paper showed that this model is able to replicate the main aspects of the dynamic behavior of inflation and output found in the data, including the persistence of the movements in these variables

A number of applications and extensions are left for future research. For instance, it would be interesting to conduct a similar analysis for data for other countries and periods, and use the results to compare price and wage setting under different macroeconomic conditions. Similarly, it would be of interest to modify the base model to allow for a distribution of price durations that varies through time, or as a function of average inflation and other aggregate variables. Another relevant extension would be to integrate the model presented in this paper with an appropriate model for the determination of the prices of farm and imported goods, for which the basic model presented in Section II seems less adequate. Finally, and perhaps most important, also left for future research is the exploration of the implications of the model estimated in this paper when embedded in complete quantitative macroeconomic or dynamic-general-equilibrium models, including its consequences for the econometric evaluation of competing monetary policies.

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