The Timing of Monetary Policy Shocks

By Giovanni Olivei and Silvana Tenreyro*

A vast empirical literature has documented delayed and persistent effects of monetary policy shocks on output. We show that this finding results from the aggregation of output impulse responses that differ sharply depending on the timing of the shock. When the monetary policy shock takes place in the first two quarters of the year, the response of output is quick, sizable, and dies out at a relatively fast pace. In contrast, output responds very little when the shock takes place in the third or fourth quarter. We propose a potential explanation for the differential responses based on uneven staggering of wage contracts across quarters. Using a dynamic general equilibrium model, we show that a realistic amount of uneven staggering can generate differences in output responses quantitatively similar to those found in the data. (*JEL* E23, E24, E58, J41)

An important branch of the macroeconomics literature is motivated by the questions of whether, to what extent, and why monetary policy matters. As concerns the first two questions, substantial empirical work has led to a broad consensus that monetary shocks do have real effects on output. Moreover, the output response is persistent and occurs with considerable delay: the typical impulse response has output peaking six to eight quarters after a monetary policy shock (see, for example, Lawrence Christiano, Martin Eichenbaum, and Charles Evans 1999). As for the third question, a large class of theories points to the existence of contractual rigidities to explain why monetary policy might cause real effects on output. Theoretical models usually posit some form of nominal or real rigidity in wages or prices that is constant over time. For example, wage contracts are assumed to be staggered uniformly over time or subject to change with a constant probability at each point in time (John B. Taylor 1980; Guillermo Calvo 1983).[1]

This convenient simplification, however, may not be a reasonable approximation of reality. As a consequence of organizational and strategic motives, wage contract renegotiations may occur at specific times in the calendar year. While there is no systematic information on the timing of wage contracts, anecdotal evidence supports the notion of "lumping" or uneven staggering of contracts. For example, evidence from firms in manufacturing, defense, information technology, insurance, and retail in New England surveyed by the Federal Reserve System in 2003 for the "Beige Book" indicates that most firms take decisions regarding compensation changes (base pay and health insurance) during the fourth quarter of the calendar year. Changes in compensation then become effective at the very beginning of the next year. The Radford Surveys of compensation practices in the information technology sector reveal that more than 90 percent of companies use a focal base-pay ad-

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[1] State-dependent versions of price- and wage-setting behavior have been developed in the literature (see Michael Dotsey, Robert King, and Alexander Wolman 1999). As we argue in the text, however, the probability of changing prices and wages over time may change for reasons not captured by changes in the state of the economy.

ministration, with annual pay-change reviews; pay changes usually take place at the beginning of the new fiscal year. According to the same surveys, 60 percent of the information technology companies close their fiscal year in December.[2] To the extent that there is a link between pay changes and the end of the fiscal year, it is worth noting that 64 percent of the firms in the Russel 3,000 Index end their fiscal year in the fourth quarter, 16 percent in the first, 7 percent in the second, and 13 percent in the third.[3] Finally, reports on collective bargaining activity compiled by the Bureau of Labor Statistics show that the distribution of expirations and wage reopening dates tends to be tilted toward the second semester of the year.[4]

If the staggering of wage contracts is not uniform, as the anecdotal evidence suggests, in principle monetary policy can have different effects on real activity at different points in time. Specifically, monetary policy should have, other things equal, a smaller impact in periods of lower rigidity—that is, when wages are being reset. This paper provides an indirect test for the presence and the importance of the lumping or uneven staggering of contracts, by examining the effect of monetary policy shocks at different times in the calendar year. In order to do so, we introduce quarter-dependence in an otherwise standard VAR model. Our goal is to assess whether the effect of a monetary policy shock differs according to the quarter in which the shock occurs and, if so, whether this difference can be reconciled with uneven staggering.

We find that there are significant differences in output impulse responses depending on the timing of the shock. In particular, after a monetary shock that takes place in the first quarter, the response of output is fairly rapid, with output reaching a level close to the peak effect four quarters after the shock. The response is even more front-loaded and dies out faster when the shock takes place in the second quarter. In this case, the peak effect is attained three quarters after the shock. In both cases, the response of output to a monetary policy shock is economically relevant. An expansionary shock in either the first or the second quarter with an impact effect on the federal funds rate of -25 basis points raises output in the following 8 quarters by an average of about 25 basis points. In contrast, the response of output to a monetary shock occurring in the second half of the calendar year is small, both from a statistical and from an economic standpoint. A 25-basis-points unexpected monetary expansion in either the third or fourth quarter raises output in the 8 quarters following the shock by less than 10 basis points on average, with the effect not statistically different from zero at standard confidence levels. The well-known finding that output takes a long time to respond and is quite persistent may be interpreted as the combination of these sharply different quarterly responses.

The dynamics of output in response to a monetary policy shock at different times of the year is mirrored by the dynamics of prices and nominal wages. The price and nominal wage responses are delayed when the shock occurs in the first half of the year, whereas prices and nominal wages respond more quickly when the shock occurs in the second half of the year.

The anecdotal evidence on wage-setting practices provides an explanation for the qualitative differences in the quarterly impulse responses. It is important to gauge, however, whether uneven staggering can also explain the quantitative differences in the estimated responses. To address this issue, we calibrate a variant of the stochastic dynamic general equilibrium model presented by Christiano, Eichenbaum, and Evans (2005). The crucial modification to the setup is that we allow the probability of resetting wages to differ across quarters. We show that a realistic amount of uneven staggering in the model can quantitatively match the quarter-dependent impulse responses estimated on actual data.

Our findings also speak to the long-standing issue of whether wages play an allocative role in the economy. The larger response of economic activity following a monetary policy shock in periods when wages are more rigid suggests that wages are allocative even in the short run.

The remainder of the paper is organized as follows. Section I presents the empirical method and introduces the data. Section II presents the dynamic effects of monetary policy on different

[2] We thank Andy Rosen of Aon Consulting's Radford Surveys for providing us with the information.

[3] This information is for the year 2003 and is available from Standard and Poor's COMPUSTAT. To compute the percentages, we have weighted each firm in the Russel 3,000 Index by the firm's number of employees.

[4] See *Current Wage Developments,* various issues, Bureau of Labor Statistics.

macroeconomic aggregates and performs a set of robustness tests. Section III presents the theoretical model and confronts it with the empirical findings. Section IV offers some concluding remarks.

I. Method

A. *Empirical Model*

Our empirical analysis for measuring the effect of monetary policy shocks relies on a very general linear dynamic model of the macroeconomy whose structure is given by the following system of equations:

$$(1) \quad \mathbf{Y}_t = \sum_{s=0}^{S} \mathbf{B}(q_t)_s \mathbf{Y}_{t-s} + \sum_{s=1}^{S} \mathbf{C}(q_t)_s p_{t-s} + \mathbf{A}^y(q_t)\mathbf{v}_t^y;$$

$$(2) \quad p_t = \sum_{s=0}^{S} \mathbf{D}_s \mathbf{Y}_{t-s} + \sum_{s=1}^{S} \mathbf{g}_s p_{t-s} + v_t^p.$$

Boldface letters are used to indicate vectors or matrices of variables or coefficients. In particular, $\mathbf{Y}_t$ is a vector of nonpolicy macroeconomic variables (e.g., output, prices, and wages), and p_t is the scalar variable that summarizes the policy stance. We take the federal funds rate as our measure of policy, and use innovations in the federal funds rate as a measure of monetary policy shocks. Federal Reserve operating procedures have varied over the past 40 years, but several authors have argued that funds-rate targeting provides a good description of Federal Reserve policy over most of the period (see Ben Bernanke and Alan Blinder 1992; and Bernanke and Ilian Mihov 1998). Equation (1) allows the nonpolicy variables $\mathbf{Y}_t$ to depend on both current and lagged values of $\mathbf{Y}$, on lagged values of p, and on a vector of uncorrelated disturbances $\mathbf{v}^y$.[5] Equation (2) states that the policy variable p_t depends on both current and lagged values of $\mathbf{Y}$, on lagged values of p, and on the monetary policy shock v^p.[6] As such, the system embeds the key restriction for identifying the dynamic effects of exogenous policy shocks on the various macro variables $\mathbf{Y}$: policy shocks do not affect macro variables within the current period. Although debatable, this identifying assumption is standard in several recent VAR analyses.[7]

The model in equations (1) and (2) replicates the specification of Bernanke and Blinder (1992), with the crucial difference that we allow for time-dependence in the coefficients. Specifically, $\mathbf{B}(q_t)_s$ and $\mathbf{C}(q_t)_s$ are coefficient matrices whose elements, the coefficients at each lag, are allowed to depend on the quarter q_t that indexes the dependent variable, where $q_t = j$ if t corresponds to the j^{th} quarter of the year. The systematic response of policy takes the time-dependence feature of the nonpolicy variables into account: substituting (1) into (2) shows that the coefficients in the policy equation are indirectly indexed by q_t through their impact on the nonpolicy variables, $\mathbf{Y}_t$.[8]

Given the identifying assumption that policy shocks do not affect macro variables within the current period, we can rewrite the system in a standard VAR reduced form, with only lagged variables on the right-hand side:

$$(3) \quad \mathbf{X}_t = \mathbf{F}(L, q)\mathbf{X}_{t-1} + \mathbf{U}_t,$$

where $\mathbf{X}_t = [\mathbf{Y}_t, p(t)]'$, $\mathbf{U}_t$ is the corresponding vector of reduced-form residuals, and $\mathbf{F}(L, q)$ is a four-quarter distributed lag matrix of coefficients that allows for the coefficients at each lag to depend on the particular quarter q indexing the dependent variable. The system can then be

[5] Note that the vector of disturbances $\mathbf{v}^y$, composed of uncorrelated elements, is premultiplied by the matrix $\mathbf{A}^y(q)$ to indicate that each element of $\mathbf{v}^y$ can enter into any of the nonpolicy equations. This renders the assumption of uncorrelated disturbances unrestrictive.

[6] Policy shocks are assumed to be uncorrelated with the elements of $\mathbf{v}^y$. Independence from contemporaneous economic conditions is considered part of the definition of an exogenous policy shock. The standard interpretation of v^p is a combination of various random factors that might affect policy decisions, including data errors and revisions, preferences of participants at the FOMC meetings, politics, etc. (see Bernanke and Mihov 1998).

[7] See, among others, Bernanke and Blinder (1992), Bernanke and Mihov (1998), Christiano et al. (1999, 2005), Jean Boivin and Marc Giannoni (2006), and Julio Rotemberg and Michael Woodford (1997).

[8] In this specification, the coefficients $\mathbf{D}_s$ and $\mathbf{g}_s$ are constant across seasons, neglecting differential policy responses in different seasons beyond the indirect effect through $\mathbf{Y}_t$ we already mentioned. We are unaware of any evidence suggesting that policy responses to given outcomes vary by season.

estimated equation by equation using ordinary least squares. The effect of policy innovations on the nonpolicy variables is identified with the impulse-response function of $\mathbf{Y}$ to past changes in v^p in the unrestricted VAR (3), with the federal funds rate placed last in the ordering.[9] An estimated series for the policy shock can be obtained via a Choleski decomposition of the covariance matrix of the reduced-form residuals.

One important implication of quarter dependence is that the effects of monetary policy shocks vary depending on the quarter in which the shock takes place. Denote by $\mathbf{X}(T)$ the skip-sampled matrix series, with $\mathbf{X}(T) = (\mathbf{X}_{1,T}, \mathbf{X}_{2,T}, \mathbf{X}_{3,T}, \mathbf{X}_{4,T})$, where $\mathbf{X}_{j,T}$ is the vector of variables in quarter j in year T, and $j = 1, 2, 3, 4$.[10] Then we can rewrite the quarter-dependent reduced-form VAR (3) as follows:

$$\text{(4)} \qquad \Xi_0 \mathbf{X}(T) = \Xi_1(L)\mathbf{X}(T-1) + \mathbf{U}(T),$$

where Ξ_0 and $\Xi_1(L)$ are parameter matrices containing the parameters in $\mathbf{F}(L, q)$ in (3), and $\mathbf{U}(T) = (\mathbf{U}_{1,T}, \mathbf{U}_{2,T}, \mathbf{U}_{3,T}, \mathbf{U}_{4,T})$, with $\mathbf{U}_{j,T}$ the vector of reduced-form residuals in quarter j of year T. The system in (4) is simply the reduced-form VAR (3) rewritten for annually observed time series. As such, the reduced-form (4) does not contain time-varying parameters. Moreover, because the matrix Ξ_0 can be shown to be lower-block-triangular, it can be inverted to yield:

$$\text{(5)} \qquad \mathbf{X}(T) = \Xi_0^{-1}\Xi_1(L)\mathbf{X}(T-1) + \Xi_0^{-1}\mathbf{U}(T),$$

with Ξ_0^{-1} still being a lower bloc-triangular matrix. The system (5) illustrates that when a monetary policy shock occurs in the first quarter, the response of the nonpolicy variables in the next quarter will be governed by the reduced-form dynamics of the nonpolicy variables in the second quarter. The response two quarters after the initial shock will be governed by the reduced-form dynamics of the nonpolicy variables in the third quarter, and so on.

B. *Testing*

The quarter-dependent VAR in (3) generates four different sets of impulse responses to a monetary policy shock, according to the quarter in which the shock occurs. It is then important to assess whether the quarter-dependent impulse-response functions are statistically different from the impulse responses of the nested standard VAR with no time dependence. A first natural test for the empirical relevance of quarterly effects consists of simply comparing the estimates obtained from the quarter-dependent VAR (3) with those obtained from the restricted standard VAR using an F-test, equation by equation. A rejection of the null hypothesis of no seasonal dependence would imply that the system generates four different sets of impulse responses. The F-tests on the linear reduced-form VAR, however, do not map one for one into a test on the corresponding impulse responses because the impulse-response functions are nonlinear combinations of the estimated coefficients in the VAR. To assess the significance of quarter-dependence directly on the impulse-response functions, we develop a second test that complements the F-test on the linear VAR equations. Specifically, we consider the maximum difference, in absolute value, between the impulse responses of variable x in the quarter-dependent VAR and in the standard non-time-dependent VAR, to obtain the following statistic:

$$D = \sup_k |x_k^q - x_k|,$$

where x_k^q denotes the period k response in the quarter-dependent model and x_k the response in the standard non-time-dependent model.[11] We resort to simulation methods for inference. Using a bootstrap procedure, we calculate the distribution of the D statistic under the assumption that there is no quarter-dependence. The bootstrap algorithm involves generating a random sample by sampling (with replacement) from

[9] The ordering of the variables in $\mathbf{Y}_t$ is irrelevant. Since identification of the dynamic effects of exogenous policy shocks on the macro variables $\mathbf{Y}$ requires only that policy shocks not affect the given macro variables within the current period, it is not necessary to identify the entire structure of the model.

[10] If $t = 1, \ldots, n$, then the observations in $\mathbf{X}_{1,T}$ are given by $t = 1, 5, 9, \ldots, n-3$, the observations in $\mathbf{X}_{2,T}$ are given by $t = 2, 6, 10, \ldots, n-2$, and so on.

[11] We compute the supremum of the difference in impulse-response functions over 20 quarters following a monetary policy shock.

the residuals of the estimated non-time-dependent reduced-form VAR. Using fixed initial conditions,[12] we recursively generate a new dataset using the estimated parameters from the standard non-time-dependent VAR. We then estimate new impulse responses from both the quarter-dependent and the standard VAR, and compute a new value D^S, where the superscript s denotes a simulated value. The procedure is repeated 2,000 times to obtain a bootstrap *p*-value, which is the percentage of simulated D^S exceeding the observed D.

C. *Data and Estimation*

Our benchmark analysis is based on quarterly data covering the period 1966:Q1 to 2002:Q4. The beginning of the estimation period is dictated by the behavior of monetary policy. Only after 1965 did the federal funds rate, the policy variable in our study, exceed the discount rate and hence act as the primary instrument of monetary policy. We use seasonally adjusted data, but in the robustness section we also present results based on non–seasonally adjusted data. The nonpolicy variables in the system include real GDP, the GDP deflator, and an index of spot commodity prices.[13] As is now standard in the literature, the inclusion of the commodity price index in the system is aimed at mitigating the "price puzzle," whereby a monetary tightening initially leads to a rising rather than falling price level. In Sections IIC and IID we discuss alternative empirical specifications that replace some of the nonpolicy variables from the baseline specification with different variables and/or include an expanded set of nonpolicy variables. The additional nonpolicy variables entering these specifications are the core Consumer Price Index (CPI), an index for wages given by compensation per hour in the nonfinancial corporate sector, and an index of hours of production workers in the manufacturing sector.[14]

We estimate each equation in the VAR (3) separately by OLS, using four lags of each variable in the system. In our benchmark specification, all the variables in the vector $\mathbf{Y}$ are expressed in log levels. The policy variable, the federal funds rate, is expressed in levels. We formalize trends in the nonpolicy variables as deterministic, and allow for a linear trend in each of the equations of the VAR (3). In the robustness section we discuss findings when GDP is expressed as the (log) deviation from a segmented deterministic trend, while the GDP deflator and the commodity price index are expressed in (log) first-differences.

II. The Dynamic Effects of Monetary Policy Shocks

A. *Results from the VAR Specification*

In this section we present the estimated dynamic effects of monetary policy shocks on real GDP, the GDP deflator, and the federal funds rate. Impulse responses are depicted in Figures 1 through 5, together with 95 percent and 80 percent confidence bands around the estimated responses.[15] We consider a monetary policy shock that corresponds to a 25-basis-point decline in the funds rate on impact. For ease of comparison, the response of the variables to the shock is graphed on the same scale across figures. Figure 1 displays impulse responses to the policy shock when we do not allow for quarter-dependence in the reduced-form VAR, as is customary in the literature. The top panel shows that the output response to the policy shock is persistent, peaking seven quarters after the shock and slowly decaying thereafter. The response of output is still more than half the peak response 12 quarters after the shock. The center panel shows that prices start to rise reliably three quarters after the shock, although it takes

[12] The fixed initial conditions are given by the values that the variables included in the VAR take over the period 1965:Q1 to 1965:Q4.

[13] The source for real GDP and the GDP deflator is the Bureau of Economic Analysis, Quarterly National Income and Product Accounts. The source for the spot commodity price index is the Commodity Research Bureau.

[14] The source for core Consumer Price Index (CPI), compensation per-hour in the nonfinancial corporate sector, and aggregate weekly hours of production workers in the manufacturing sector is the Bureau of Labor Statistics.

[15] While much applied work uses 95 percent confidence intervals, Christopher Sims and Tao Zha (1999) note that the use of high-probability interval camouflages the occurrence of large errors of over-coverage and advocate the use of smaller intervals, such as intervals with 68 percent coverage (one standard error in the Gaussian case). An interval with 80 percent probability corresponds to about a standard error of 1.3 in the Gaussian case.

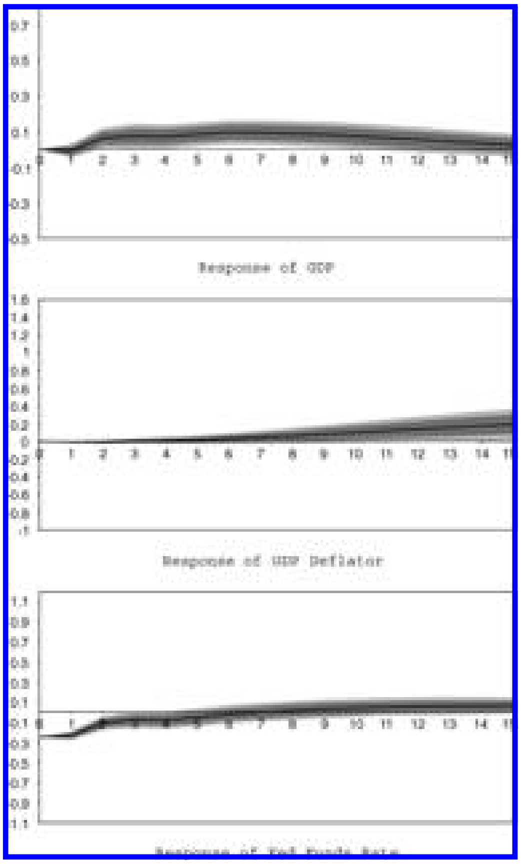

FIGURE 1. 25-BASIS-POINT DECLINE IN FED FUNDS RATE
(*No quarterly dependence: 1966:Q1 to 2002:Q4*)

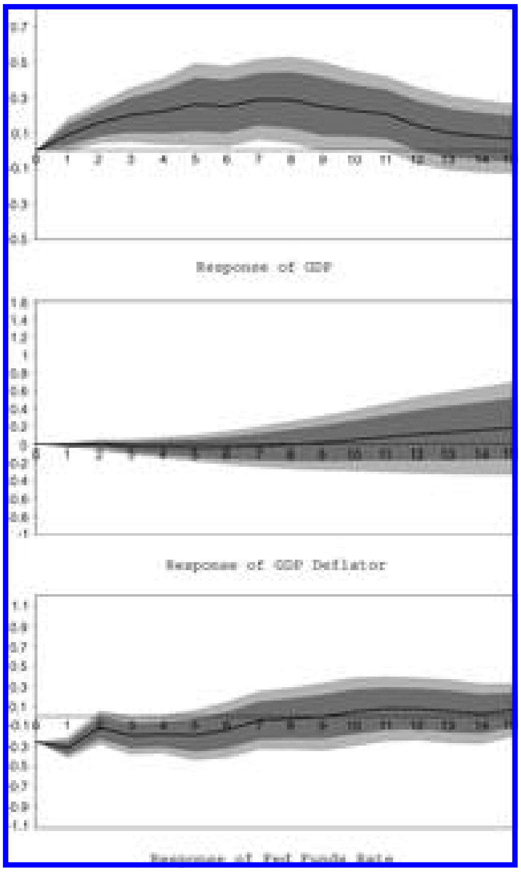

FIGURE 2. 25-BASIS-POINT DECLINE IN FED FUNDS RATE IN Q1
(*Quarterly dependence: Benchmark model 1966:Q1 to 2002:Q4*)

about one and a half years for the increase to become significant. The bottom panel, which displays the path of the federal funds rate, illustrates that the impact on the funds rate of a policy shock is less persistent than the effect on output.

Figures 2 to 5 display impulse responses when we estimate the quarter-dependent reduced-form VAR (3). The responses to a monetary policy shock occurring in the first quarter of the year are shown in Figure 2. Output rises on impact and reaches a level close to its peak response four quarters after the shock. The output response dies out at a faster pace than in the non-time-dependent VAR: 12 quarters after the shock, the response of output is less than a third of the peak response, which occurs 7 quarters after the shock, as in the non-time-dependent VAR. Moreover, the peak response is now more than twice as large as in the case with no quarter-dependence. The center panel shows that, despite controlling for commodity prices, there is still a "price puzzle," although the decline in prices is not statistically significant. It takes about seven quarters after the shock for prices to start rising. The federal funds rate, shown in the bottom panel, converges at about the same pace as in Figure 1.

Figure 3 displays impulse responses to a shock that takes place in the second quarter. It is apparent that the response of output is fast and sizable. Output reaches its peak three quarters after the shock, and the peak response is more than three times larger than the peak response in the case with no quarter dependence. Moreover, the response wanes rapidly, becoming insignificantly different from zero eight quarters after

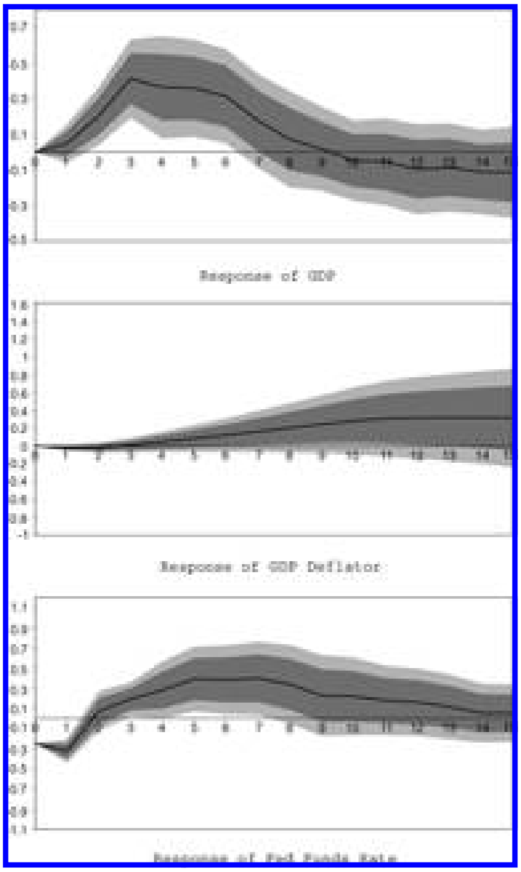

FIGURE 3. 25-BASIS-POINT DECLINE IN FED FUNDS RATE IN Q2
(*Quarterly dependence: Benchmark model 1966:Q1 to 2002:Q4*)

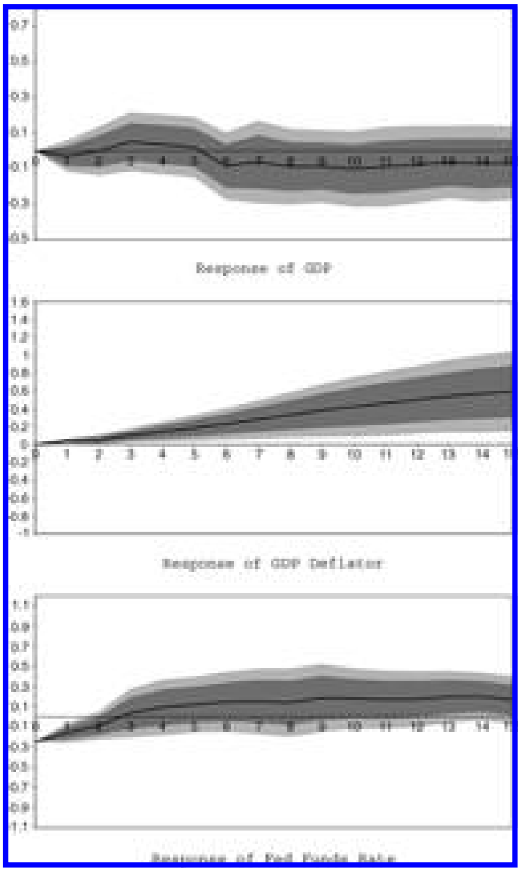

FIGURE 4. 25-BASIS-POINT DECLINE IN FED FUNDS RATE IN Q3
(*Quarterly dependence: Benchmark model 1966:Q1 to 2002:Q4*)

the shock. The center panel shows that prices start rising three quarters after the shock. The bottom panel illustrates that the large output response occurs despite the fact that the policy shock exhibits little persistence.

The responses to a monetary policy shock in the third and the fourth quarter of the year contrast sharply with the responses to a shock taking place in the first half of the calendar year. Figure 4 shows the impulse responses to a shock that occurs in the third quarter. The response of output in the top panel is now small and insignificant, both from a statistical and from an economic standpoint. Interestingly, as the center panel illustrates, prices start to increase reliably immediately after the shock. The output and price responses to a shock in the fourth quarter are qualitatively similar. As Figure 5 illustrates, the response of output is fairly weak, while prices respond almost immediately following the shock.

The differences in output responses are substantial from a policy standpoint. The policy shock raises output in the following eight quarters by an average of almost 25 basis points in either the first or the second quarter. In contrast, the increase in output is less than 10 basis points on average in both the third and the fourth quarter. Moreover, the differences documented in Figures 1 to 5 are corroborated by formal tests on the importance of quarter-dependence. Equation-by-equation *F*-tests in the reduced-form VAR (3) yield *p*-values of 0.18 for the output equation, 0.04 for the price equation, 0.03 for the commodity prices equation, and 0.006 for the federal funds rate equation.

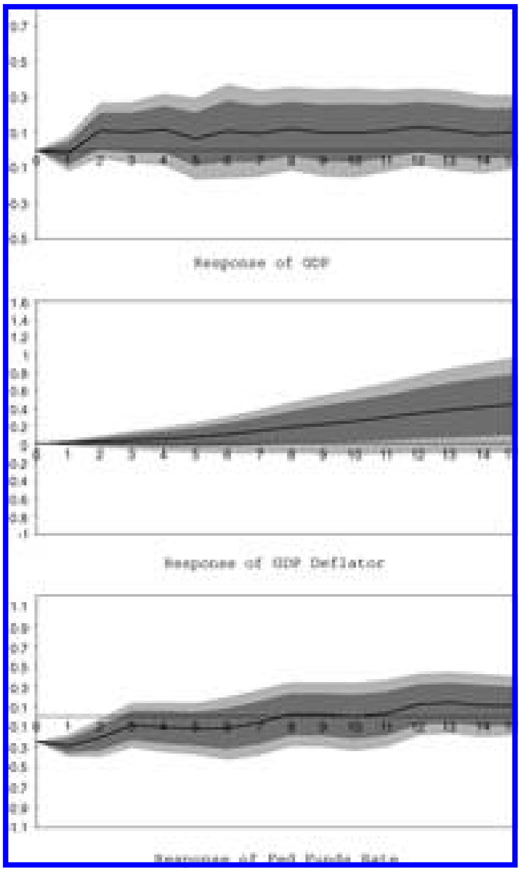

FIGURE 5. 25-BASIS-POINT DECLINE IN FED FUNDS RATE IN Q4
(*Quarterly dependence: Benchmark model 1966:Q1 to 2002:Q4*)

While indicative of the existence of seasonal dependence, these relatively low p-values do not necessarily translate into statistically different impulse responses for the corresponding variables. For this purpose, we evaluate the D-statistic described in Section IB, which assesses whether the maximum difference between the impulse response of a given variable in the quarter-dependent VAR and the corresponding response of that variable in the standard non-time-dependent VAR is statistically different. Table 1 reports the bootstrapped p-values for the D-statistic in each quarter for GDP, the GDP deflator, and the federal funds rate. The table shows that, according to this test, the output response in the first, second, and third quarter of the calendar year is statistically different from the non-time-dependent output impulse response at better than the asymptotic 5 percent level. The null hypothesis of an output response equal to the non-time-dependent response is rejected at the asymptotic 10 percent level in the fourth quarter. As may be inferred from Figures 1 through 5, the table shows that the evidence in favor of quarter-dependent price impulse responses is weaker than for output. Still, the test identifies statistical differences in the third and fourth quarters: in the third quarter, the null hypothesis of a price response equal to the non-time-dependent response is rejected at the asymptotic 5 percent level, and in the fourth quarter the null is rejected at the 10 percent level.

TABLE 1—DIFFERENCES IN IMPULSE RESPONSES ACROSS QUARTERS
(*p-values for D-Statistic*)

	Quarter			
Variable	First	Second	Third	Fourth
GDP	0.01	0.00	0.03	0.08
GDP deflator	0.39	0.16	0.02	0.06
Fed funds rate	0.22	0.00	0.12	0.02

B. *The Distribution of Monetary Policy Shocks and the State of the Economy*

We now consider whether the different impulse responses we obtain across quarters are the result of different types of shocks. In principle, differences in the direction (expansionary versus contractionary) of shocks could produce different impulse responses.[16] To explore this issue, we test for the equality of the distributions of shocks across quarters by means of a Kolmogorov-Smirnov test. The test consists of a pairwise comparison of the distributions of shocks between every two quarters, with the null hypothesis of identical distributions. We find that we cannot reject the null hypothesis in any two quarters: the smallest p-value corresponds to the test for the equality of the distributions of shocks between the third and fourth quarters and is equal to 0.31; the largest p-value corresponds to the test between the second and third quarters and is equal to 0.97. These findings suggest that differences in the direction of

[16] See Harald Uhlig (2005).

the shocks across quarters are unlikely to provide an explanation for the quarterly differences in impulse responses documented in Figures 2 through 5.

Another important issue is whether our findings are driven by the state of the economy. In principle, a theoretical argument can be made that an expansionary monetary policy shock has a larger impact on output and a smaller impact on prices when the economy is running below potential, and, vice versa, a smaller impact on output and a larger impact on prices when the economy is running above potential. To explore this issue, we partitioned the data according to whether the output gap was positive or negative and estimated two different reduced-form VARs. The impulse responses for output and prices to a monetary policy shock from the VAR estimated using observations corresponding to a negative output gap were similar to the impulse responses obtained from the VAR estimated using observations corresponding to a positive output gap.[17] These results suggest that the stage of the business cycle is unlikely to be a candidate for explaining the different impulse responses across quarters.

There is, however, a more subtle way in which the state of the economy could influence our findings. Robert Barsky and Jeffrey Miron (1989) trace a parallel between seasonal and business cycles, and note that in seasonally unadjusted data the first and third quarters resemble a recession (the third quarter being milder), whereas the second and fourth quarters resemble an expansion (the fourth being stronger). Our use of seasonally adjusted data should, in principle, control for the seasonal component of output. And even if such a control were imperfect, the pattern of impulse responses in Figures 2 to 5 cannot be easily reconciled with the seasonal cycle. The response of output is, in fact, large when the policy shock occurs in the first (recession) and second (expansion) quarters, and the response is weak when the shock occurs in the third (recession) and fourth (expansion) quarters. It is still possible, though, that the seasonal pattern of some activities that are particularly sensitive to interest rate movements, such as (consumption and producer) durables and structures, could affect some of our findings. In particular, Barsky and Miron (1989) show that in the first quarter of the year there is a pronounced seasonal slowdown in spending for both durables and structures. A potential interpretation for our fourth-quarter results would then be that at that time of the calendar year, a policy shock has little impact on output because the transmission mechanism is impaired by the low level of interest-sensitive activities in the subsequent quarter. This interpretation, however, does not fit well with our third-quarter results. As Barsky and Miron (1989) also show, the seasonal level of spending on producer and consumer durables is especially high in the fourth quarter. If the previous reasoning is correct, we should observe a large response of output to shocks taking place in the third quarter, which is in conflict with the evidence. We thus view our findings as difficult to reconcile with an explanation that relies mainly on seasonal fluctuations in output or in the interest-sensitive components of output.

C. *Robustness Checks*

We now summarize results on the robustness of our baseline specification along several dimensions.[18] Because the quarter-dependent reduced-form VAR (3) requires the estimation of a fairly large number of parameters, we investigated whether our findings are sensitive to outliers. For this purpose, we reestimated the VAR equation by equation using Peter Huber's (1981) robust procedure. The Huber estimator can be interpreted as a weighted least squares estimator that gives a weight of unity to observations with residuals smaller in absolute value than a predetermined bound, but downweights outliers (defined as observations with residuals larger than the predetermined bound).[19] With

[17] These findings are available upon request. There is no established evidence in the extant empirical literature that monetary shocks have different effects according to the stage of the business cycle.

[18] For reasons of space, we do not provide pictures of the estimated impulse responses in most cases. The pictures not shown in the paper are available from the authors upon request.

[19] Denote by σ the standard deviation of the residuals in any given equation of the VAR (3). For the given equation, the Huber estimator gives a weight of unity to observations with residuals smaller in absolute value than $c\sigma$, where c is a parameter usually chosen in the range $1 \leq c \leq 2$, while outliers, defined as observations with residuals larger than $c\sigma$, receive a weight of $(c\sigma/|u_i|)$, where u_i is the residual for

this procedure, estimated impulse responses (not shown) turn out to be very similar to the ones in Figures 2 through 5. A notable byproduct of the robust estimation procedure is that p-values for the F-tests on quarter-dependence are now well below 0.05 for all the equations in the VAR.

Since our proposed explanation for the different impulse responses relies on uneven staggering of wage contracts across quarters, we considered an alternative specification where the GDP deflator is replaced by a nominal wage index in the vector of nonpolicy variables $\mathbf{Y}$. Using wages in lieu of final prices does not alter our main findings. The estimated output responses (not shown) are virtually the same as in the benchmark specification, and the response of wages closely mimics the response of prices across different quarters.

In our benchmark specification we control for seasonal effects by using seasonally adjusted data. Still, because we are exploiting a time-dependent feature of the data, it is of interest to check whether our results are driven by the seasonal adjustment. To this end, we estimate impulse responses to a monetary shock from the quarter-dependent reduced-form VAR (3) using seasonally unadjusted data for the nonpolicy variables $\mathbf{Y}$.[20] The results from this exercise are illustrated in Figures 6 through 9, which show the responses of output, prices, and the federal funds rate to a 25-basis-point decline in the funds rate. The responses of output and prices using seasonally unadjusted data are remarkably similar to the responses obtained in the benchmark specification using seasonally adjusted data, although the estimated responses with seasonally unadjusted data are less precise.

The results continue to hold under a different treatment of the low-frequency movements in output and prices. Specifically, we considered a specification for the quarter-dependent reduced-form VAR (3) in which variables in $\mathbf{Y}$ are not expressed in log levels, but rather GDP is expressed as a deviation from a segmented deterministic trend,[21] and prices are expressed in log first-differences. Such a specification is used in several papers in the literature (see, e.g., Boivin and Giannoni 2006). The estimated impulse responses (not shown) are qualitatively similar to those reported in the benchmark specification. In particular, output responds more strongly and more quickly to a monetary policy shock in the first than in the second half of the year, while the opposite occurs for inflation.

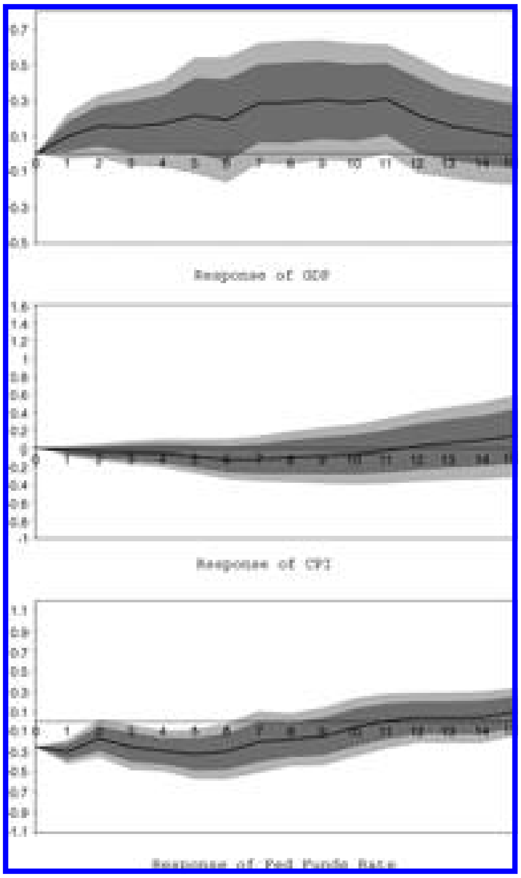

FIGURE 6. 25-BASIS-POINT DECLINE IN FED FUNDS RATE IN Q1
(*Quarterly dependence: NSA data system 1966:Q1 to 2002:Q4*)

observation i. The estimator is fairly insensitive to the choice of c.

[20] Since we do not have data on the seasonally unadjusted GDP deflator, we replace the GDP deflator with the seasonally unadjusted CPI. The CPI index is also used to deflate the seasonally unadjusted data for nominal GDP.

[21] Specifically, we consider the deviation of log real GDP from its segmented deterministic linear trend, with breakpoints in 1974 and 1995.

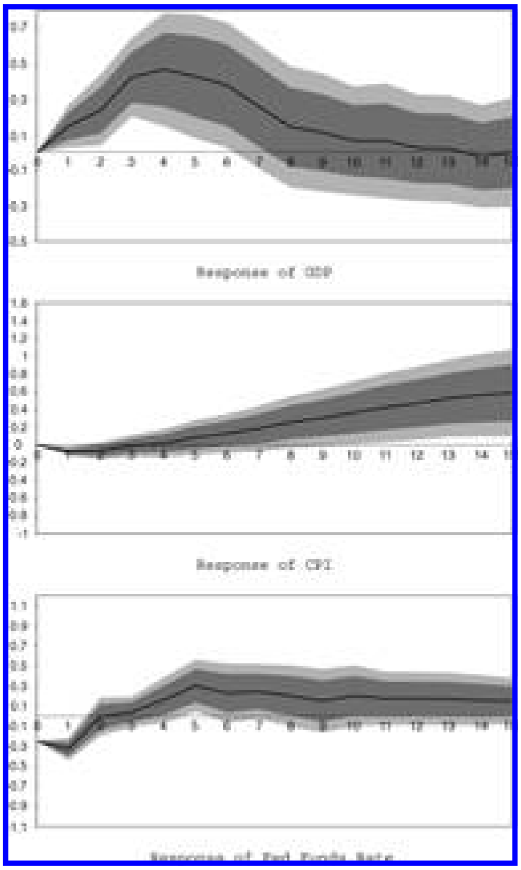

FIGURE 7. 25-BASIS-POINT DECLINE IN FED FUNDS RATE IN Q2
(*Quarterly dependence: NSA data system 1966:Q1 to 2002:Q4*)

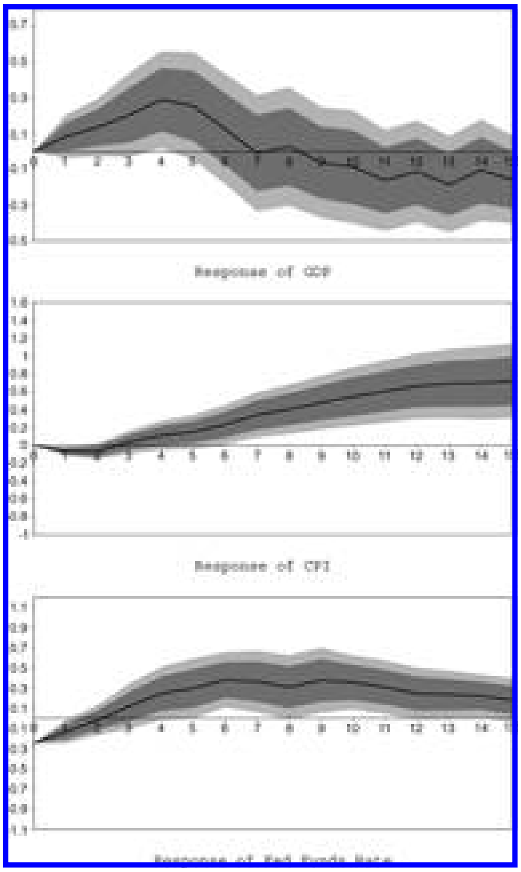

FIGURE 8. 25-BASIS-POINT DECLINE IN FED FUNDS RATE IN Q3
(*Quarterly dependence: NSA data system 1966:Q1 to 2002:Q4*)

D. *Additional Evidence and Interpretation*

Overall, Figures 2 through 5 and the supporting statistics uncover considerable differences in the response of output across quarters. The slow and persistent response of output to a policy shock typically found in the literature and reported in Figure 1 is the combination of different quarter-dependent responses. The difference in the behavior of prices across quarters is somewhat less striking, given the imprecision with which the responses are estimated. It is interesting, though, that when the policy shock occurs in the third and fourth quarters, prices rise more quickly than when the shock takes place in either the first or the second quarter.

Since our explanation for the main empirical findings hinges on uneven staggering of wage contracts, it is of interest to examine the quarterly responses of relevant labor market variables to a monetary policy shock. Specifically, we consider the responses of real wages and aggregate hours, estimated from a quarter-dependent VAR that includes real GDP, the GDP deflator, the real wage, hours, and the federal funds rate.[22] Figures 10 to 13 depict the quarter-dependent responses for the real wage and hours to a monetary policy shock.[23] We

[22] The real wage is computed as the ratio of compensation per hour in the nonfinancial corporate sector and the GDP deflator. All the nonpolicy variables are expressed in log levels.

[23] The estimated responses (not shown) for GDP, the GDP deflator, and the federal funds rate are qualitatively similar to the responses depicted in Figures 2 to 5.

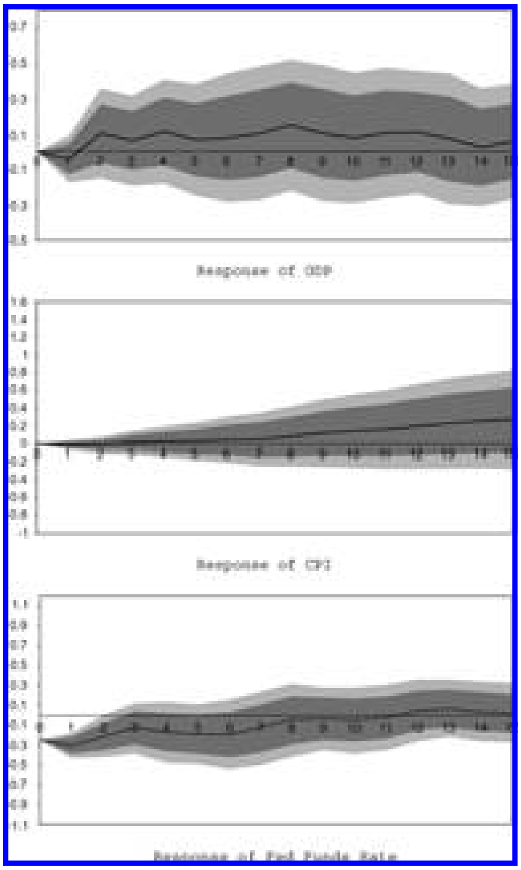

FIGURE 9. 25-BASIS-POINT DECLINE IN FED FUNDS RATE IN Q4
(*Quarterly dependence: NSA data system 1966:Q1 to 2002:Q4*)

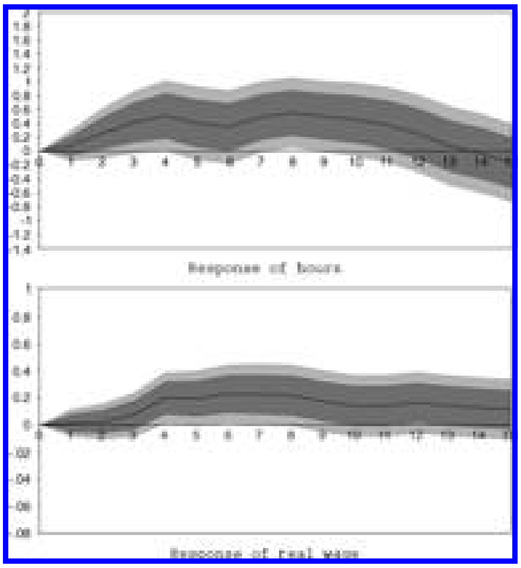

FIGURE 10. 25-BASIS-POINT DECLINE IN FED FUNDS RATE IN Q1
(*Wage-hour data system* 1966:*Q*1 *to* 2002:*Q*4)

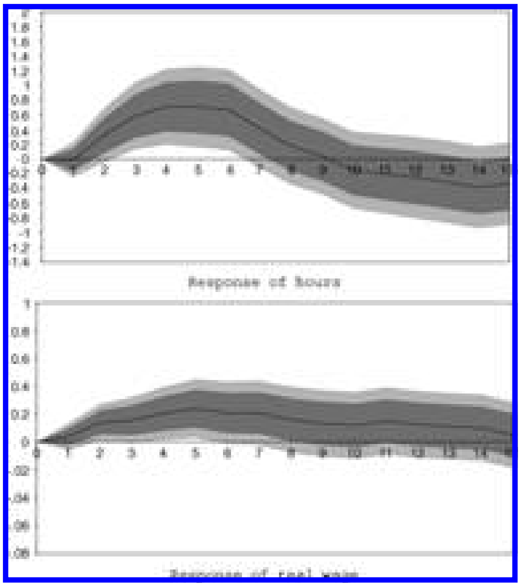

FIGURE 11. 25-BASIS-POINT DECLINE IN FED FUNDS RATE IN Q2
(*Wage-hour data system 1966:Q1 to 2002:Q4*)

mentioned in the previous section that the responses of the nominal wage are qualitatively similar to the responses of prices across different quarters. The additional information conveyed by the present exercise is that the estimated responses for the real wage are mildly procyclical and similar across quarters. In contrast, the estimated responses for hours exhibit a noticeable quarterly pattern. The pattern mimics the estimated responses for output, with larger and statistically more significant responses when the shock occurs in the first and second quarters of the calendar year. In the next section, we provide an explanation for the contrasting dynamics of hours and the real wage.

Taken together, the VAR results and the anecdotal evidence on uneven staggering lend themselves to the following interpretation. If a large number of firms sign wage contracts at the end of the calendar year, then, on average, monetary policy shocks in the first half of the year will have a large impact on output, with little effect on nominal wages and prices. In contrast,

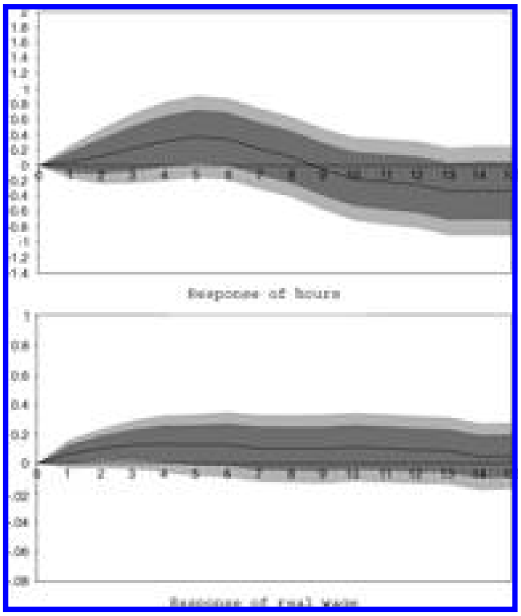

FIGURE 12. 25-BASIS-POINT DECLINE IN FED FUNDS RATE IN Q3
(*Wage-hour data system 1966:Q1 to 2002:Q4*)

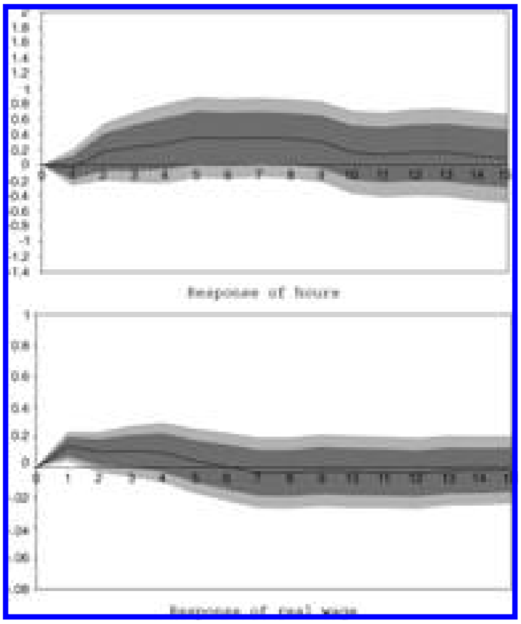

FIGURE 13. 25-BASIS-POINT DECLINE IN FED FUNDS RATE IN Q4
(*Wage-hour data system 1966:Q1 to 2002:Q4*)

monetary policy shocks in the second half of the year will be quickly followed by nominal wage and price adjustments. The policy shock will be "undone" by the new contracts at the end of the year and, as a result, the effect on output will be smaller on average. We next formalize this interpretation in the context of a dynamic stochastic general equilibrium setup that can quantitatively account for the quarter-dependent impulse responses estimated in the data.

III. A Model of Uneven Staggering

The anecdotal evidence on uneven staggering of wage-setting decisions provides an intuitively appealing explanation for the finding of quarter dependence in the response of the economy to monetary policy shocks. In this section, we investigate whether this qualitative mechanism is also quantitatively relevant. In particular, we ask whether a variant of the model proposed by Christiano et al. (2005) that allows for a realistic degree of uneven staggering of wage contracts can quantitatively match the quarter-dependent impulse responses in the data.

Christiano et al.'s (2005) setup embeds a simple form of contractual rigidities based on Calvo (1983), whereby workers face a constant probability of reoptimizing their nominal wage every quarter. Here, we generalize the setup and allow the probability of changing wages to differ across quarters in a calendar year. This modeling strategy is a simple way of introducing clustering of wage contracts at certain times of the calendar year, and it nests the standard Calvo-style sticky-wage framework as a special case when the probability of resetting wages is constant across quarters.

The model's key propagation mechanism in response to a temporary shock to the nominal interest rate is as follows. After an expansionary monetary shock, aggregate demand goes up and so does the derived demand for labor. Consider first a period when a large fraction of the workers cannot readjust their wage optimally. These workers will have to supply as much labor as the firms demand at the preset wage, leading to an increase in total hours of work (and in the usage of other inputs) and hence to an increase in aggregate activity. Consider now a period in which a large fraction of the workers can readjust their wage optimally. Following an expansionary monetary policy shock, firms can no longer count on as much labor at preset wages as in the previous scenario, as they now meet an upward-sloping supply of labor for a large number of workers. The increase in aggregate activity is therefore smaller than in the previous case, and the economy operates at a level closer to the flexible-wage-price potential.

In the model, monetary policy is neutral in the long run. Consistent with our empirical setup, monetary policy targets a short-term interest rate by means of a Taylor-type reaction function. Since interest rate targeting by the monetary authority makes money demand irrelevant in determining the equilibrium evolution of the other variables in the model, we consider a cashless version of the model.[24] As is now standard in the literature, the model features habit formation in consumption. In addition, the model incorporates investment adjustment costs. The motivation for adjustment costs to consumption and investment is largely empirical. These costs attenuate the initial impact of a monetary policy shock on the economy, and make the effects of the shock long-lasting.

In what follows, we model the behavior of firms and households and the evolution of nominal wages in the presence of uneven staggering. We then calibrate the model parameters and compare the model-generated impulse responses with the empirical responses from our quarter-dependent benchmark VAR specification and with the additional responses reported in Section IID.

A. *Firms*

The final consumption good in the economy, Y_t, is produced by a perfectly competitive firm that uses a continuum of intermediate goods, Y_{jt}, combined through the CES technology:

$$Y_t = \left(\int Y_{jt}^{1/\lambda_f} \, dj \right)^{\lambda_f},$$

with $\lambda_f \in [1, \infty)$. The firm takes the final consumption good price, P_t, and the intermediate good price, P_{jt}, as given. Each intermediate good $j \in (0, 1)$ is produced by a monopolist firm using the technology:

$$Y_{jt} = k_{jt}^{\alpha} L_{jt}^{1-\alpha} - \phi, \alpha \in (0, 1) \qquad \text{if } k_{jt}^{\alpha} L_{jt}^{1-\alpha} \geq \phi,$$

$$= 0 \qquad \text{otherwise,}$$

where L_{jt} and k_{jt} are the labor and capital services used in the production of intermediate j, and ϕ represents a fixed cost of production. Intermediate good firms rent capital at the rental rate R_t^k and labor at the average wage rate W_t in perfectly competitive factor markets. Profits are distributed to households at the end of each period. Workers are paid in advance of production, hence the jth firm borrows its wage bill $W_t L_{jt}$ from a financial intermediary at the beginning of the period, repaying at the end of the period at the gross interest rate R_t.

Intermediate firm j sets price $P_{j,t}$ following a Calvo-type rule, with a probability $(1 - \xi_p)$ of being able to reoptimize its nominal price in each period. The ability to reoptimize is independent across firms and time. Firms that cannot reoptimize index their price to lagged inflation: $P_{j,t} = \pi_{t-1} P_{j,t-1}$, where $\pi_{t-1} = P_t / P_{t-1}$.

B. *Households*

There is a continuum of infinitely lived households i, with $i \in (0, 1)$, which derive utility from consumption and disutility from labor effort, and which own the economy's stock of physical capital. Every period each household makes the following sequence of decisions. First, the household decides how much to consume and invest in physical capital, and how much capital services to supply. Second, the household purchases securities that are contingent on whether it can reoptimize its wage.[25] Third, the household sets its wage after finding

[24] See Michael Woodford (2003).

[25] As in Christiano et al. (2005), this is assumed for analytical convenience. Because the uncertainty over whether a household can reoptimize wages is idiosyncratic, households work different numbers of hours and earn different wage rates; this, in turn, could lead to potentially heterogeneous levels of consumption and asset holdings. In equilibrium, the existence of state-contingent securities ensures that households are homogeneous with respect to consumption choice and asset holdings, though heterogeneous with respect to the wage rate and the supply of labor.

out whether it can reoptimize. Household i's expected utility is given by[26]

(6)

$$E_{t-1}^{i}\left\{\sum_{j=0}^{\infty}\beta^{t+j}[u(c_{t+j}-be_{t+j})-z(h_{i,t+j})]\right\},$$

where $\beta \in (0, 1)$ is the intertemporal discount factor; c_t is consumption at time t; u is the utility derived from consumption, with $u' > 0$ and $u'' < 0$; $e_{t+j} = \chi e_{t+j-1} + (1-\chi)c_{t+j-1}$ represents habit formation in consumption, with b, $\chi \in [0, 1)$;[27] $h_{i,t}$ is the number of hours worked at t; and z is the disutility of labor effort, with $z' > 0$ and $z'' \geq 0$.[28]

The household's dynamic budget constraint is given by

$$V_{t+1} + P_t[c_t + i_t + a(u_t)\bar{k}_t] = R_t V_t + R_t^k u_t \bar{k}_t + D_t + A_{i,t} + W_{i,t}h_{i,t} - T_t, \qquad \forall\, t,$$

where V_t is the household's beginning of period t financial wealth, deposited at a financial intermediary where it earns the gross nominal interest rate R_t. The term $R_t^k u_t \bar{k}_t$ represents the household's earnings from supplying capital services, where $u_t\bar{k}_t$ denotes the physical stock of capital $\bar{k}_t$, adjusted by the capital utilization rate u_t, whose level is decided by the household at time t, and R_t^k denotes the corresponding return. D_t denotes firm profits, and $A_{i,t}$ denotes the net cash inflow from participating in state-contingent securities markets at time t. $W_{i,t}h_{i,t}$ is labor income and T_t is a lump-sum tax used to finance government expenditures.[29] Finally, i_t and $a(u_t)\bar{k}_t$ represent, respectively, the purchases of investment goods and the cost of setting the utilization rate to u_t, in units of consumption goods.

The evolution of the stock of physical capital is given by

$$\bar{k}_{t+1} = (1-\delta)\bar{k}_t + F(i_t, i_{t-1}), \tag{7}$$

where δ is the depreciation rate, and the adjustment costs function F represents the technology transforming current and past investment into installed capital for use in the following period.[30] Capital services k_t are related to the stock of capital $\bar{k}_t$ through $k_t = u_t\bar{k}_t$.

Financial intermediaries receive V_t from households and lend all their money to intermediate-good firms, which use the money to buy labor services L_t. Loan market clearing hence implies $W_tL_t = V_t$.

C. *Wage Setting with Nominal Rigidities and Staggering*

The household enjoys monopoly power over its differentiated labor service $h_{i,t}$, and sells this service to a representative competitive firm that transforms the service into an aggregate labor input L_t, using a CES technology with parameter $\lambda_w \in [1, \infty)$. The implied demand schedule for $h_{i,t}$ is given by:

$$h_{i,t} = \left(\frac{W_t}{W_{i,t}}\right)^{\lambda_w/(\lambda_w-1)} L_t,$$

where W_t is the aggregate wage level (the price of L_t), which takes the form

$$W_t = \left[\int_0^1 W_{i,t}^{1/(1-\lambda_w)}\, di\right]^{1-\lambda_w}.$$

Households take W_t and L_t as given, and set wages knowing that they will have to supply as much labor as firms demand at the set

[26] As a consequence of the model's assumptions, the formula anticipates the result that all households consume the same, despite working different hours, hence, the corresponding omission of the subscript i in c_t.

[27] See Jeffrey Fuhrer (2000) for a thorough explanation of habit formation in consumption as represented in this functional form, and the technical appendix in Christiano et al. (2005).

[28] In the model's calibration, we assume $u(\cdot) = \log(\cdot)$ and $z(\cdot) = \vartheta_0(\cdot)^2$.

[29] We assume that the government budget is always balanced, that is, $T_t = P_t g_t$, where g_t denotes real government expenditures on the final good. In the present setup, government consumption does not substitute for households' private consumption; we assume g_t is set at a level proportional to the previous year's quarterly average GDP.

[30] In calibrating the model, we assume, as in Christiano et al., that $F(i_t, i_{t-1}) = [1 - S(i_t/i_{t-1})]i_t$, with $S'(1) = 0$, $\kappa \equiv S''(1) > 0$, $u_t = 1$ in steady state, and $a(1) = 0$. We denote $\sigma_a = a''(1)/a'(1)$.

wage. This way of modeling the labor market implies that wages play an allocative role at any point in time. We assume that wages are set following the mechanism proposed by Calvo (1983). However, instead of facing a constant probability of resetting the wage in any given period, households face different probabilities over the course of a calendar year. To keep a close parallel with the empirical exercise, we divide the year into four quarters and assume that the probability of resetting wages in quarter k is $(1 - \alpha_k)$, with $k = 1, \dots, 4$. When a household receives the signal to change its wage, the new wage is set optimally by taking into account the probability of future wage changes. In particular, if a contract is negotiated in the first quarter, the probability that it is *not* renegotiated in the second quarter will be α_2; the probability that it is *not* renegotiated in the next two quarters will be $\alpha_2\alpha_3$. Subsequent probabilities will be, correspondingly, $(\alpha_2\alpha_3\alpha_4)$, $(\alpha_2\alpha_3\alpha_4\alpha_1)$, $(\alpha_2^2\alpha_3\alpha_4\alpha_1)$, and so on. If a household cannot reoptimize its wage at time t, it uses a simple automatic rule of the form $W_{i,t} = \pi_{t-1}W_{i,t-1}$. The optimal wage for labor-type i resetting the contract in the first quarter is then

$$\tilde{W}_{i,t} = \arg\max_{\{W_{i,t}\}} E_{t-1} \tag{8}$$

$$\left\{\sum_{j=0}^{\infty}\prod_{k=1}^{4}\alpha_k^j\beta^{4j}\left[\left(\frac{\psi_{t+4j}X_{t,4j}W_{i,t}h_{i,t+4j}}{P_{t+4j}} - z(h_{i,t+4j})\right)\right.\right.$$

$$+ \alpha_2\beta\left(\frac{\psi_{t+4j+1}X_{t,4j+1}W_{i,t}h_{i,t+4j+1}}{P_{t+4j+1}} - z(h_{i,t+4j+1})\right)$$

$$+ \alpha_2\alpha_3\beta^2\left(\frac{\psi_{t+4j+2}X_{t,4j+2}W_{i,t}h_{i,t+4j+2}}{P_{t+4j+2}} - z(h_{i,t+4j+2})\right)$$

$$+ \alpha_2\alpha_3\alpha_4\beta^3\left(\frac{\psi_{t+4j+3}X_{t,4j+3}W_{i,t}h_{i,t+4j+3}}{P_{t+4j+3}}\right.$$

$$\left.\left.\left. - z(h_{i,t+4j+3})\right)\right]\right\},$$

where $X_{t,l}$ is

$$X_{t,l} = \pi_t \cdot \pi_{t+1} \cdot \pi_{t+2} \cdot \cdots \cdot \pi_{t+l-1} \quad \text{if } l \geq 0, \tag{9}$$

$$= 1 \qquad \text{if } l = 0,$$

and ψ_t is the marginal utility of (real) income at time t. The optimal wage in the first quarter maximizes the expected stream of discounted utility from the new wage, defined as the difference between the (real) gain derived from the hours worked at the new wage, $(X_{t,t'}W_{i,t}h_{i,t'}/P_{t'})$ (expressed in *utils*) and the disutility $z(h_{t'})$ of working, for all $t' \geq t$.[31] This expression is valid only for wages set in the first quarter; expressions for the optimal wage in the other quarters of the calendar year can be derived in a similar fashion.[32] Note that, since all labor types resetting their wages at a given quarter will choose the same wage, we can drop the index i and simply refer to the optimal wage in period t as $\tilde{W}_t$.

Defining variables with a hat as $\hat{x}_t = (x_t - x)/x$, where x is the value of x_t in the nonstochastic steady state, we define $\hat{\tilde{w}}_t \equiv \hat{\tilde{W}}_t - \hat{P}_t$ and $\hat{w}_t \equiv \hat{W}_t - \hat{P}_t$. Introducing the dummy variables q_{kt}, with $k = 1, \cdots 4$, which take the value of one in quarter k and zero otherwise, we show in the Appendix that wage maximization in a given quarter leads to the following recursive log-linearized expression for the optimal wage rate:

$$\hat{\tilde{w}}_t + \hat{\pi}_t = \frac{1 - \prod_{k=1}^{4}\alpha_k\beta^4}{\Gamma_{(q,\alpha,\beta)}} \tag{10}$$

$$\times\left[\hat{\pi}_t + \frac{\lambda_w - 1}{2\lambda_w - 1}\right.$$

$$\left.\times\left(\frac{\lambda_w}{\lambda_w - 1}\hat{w}_t + \hat{L}_t - \hat{\psi}_t\right)\right]$$

$$+ \beta\sum_{k=1}^{4} q_{kt}\alpha_{k+1}$$

$$\times\frac{\Phi_{(q,\alpha,\beta)}}{\Gamma_{(q,\alpha,\beta)}}E_{t-1}(\hat{\tilde{w}}_{t+1} + \hat{\pi}_{t+1}),$$

[31] Labor income is expressed in real terms (by dividing by the price level) and converted into utils using the marginal utility of income ψ.

[32] The corresponding expression for workers setting wages in the second quarter can be obtained by substituting α_{k+1} for α_k, for $k = 1, 2, 3$ and substituting α_4 for α_1 in expression (8). Similar substitutions lead to the formulae for the optimal wage in the third and fourth quarters.

with the expressions for the parameters $\Gamma_{(q,\alpha,\beta)}$ and $\Phi_{(q,\alpha,\beta)}$ provided in the Appendix. According to this expression, the optimal (real) wage is an average between a constant markup λ_w over the marginal rate of substitution between consumption and leisure, and the optimal wage expected to prevail in the next period. The weights on the two terms vary according to the quarter in which the wage is reset. In particular, the weight on the next-period wage is larger the smaller the probability of changing wages in the next period. The dynamics for the aggregate wage level W_t is given by the weighted average between the optimal wage of the labor types that receive the signal to change and the wages of those who did not get the signal. By the law of large numbers, the proportion of labor types renegotiating wages at quarter k will be equal to $(1 - \alpha_k)$. Therefore, as shown in the Appendix, the quarter-dependent law of motion for the aggregate real wage in log-deviations from the steady state is given by

$$(11)\quad \hat{w}_t = \left[1 - \left(\sum_{k=1}^{4} q_{kt}\alpha_k\right)\right]\hat{\hat{w}}_t + \left(\sum_{k=1}^{4} q_{kt}\alpha_k\right)(\hat{w}_{t-1} + \hat{\pi}_{t-1} - \hat{\pi}_t).$$

D. *Model Solution and Calibration*

The model's solution procedure involves taking a linear approximation about the nonstochastic steady state of the economy. The log-linear relationships characterizing the dynamics of the nonpolicy variables are given by equations (17) through (26) in the Appendix, together with the wage equations (10) and (11) in the text. To close the model, we assume that the monetary authority follows a Taylor-type reaction function,

$$(12)\quad \hat{R}_t = \rho\hat{R}_{t-1} + (1-\rho)(a_\pi\hat{\pi}_t + a_y\hat{y}) + \varepsilon_t,$$

where $\hat{R}_t = \ln R_t - \ln \bar{R}$ is the nominal interest rate in deviation from its steady-state level, and $\hat{y}_t$ is the output gap. The term ε_t is the policy shock, whose effect on the economy we want to evaluate. Because of the presence of the time-varying indicators q_{kt} in the equations describing the wage dynamics, the system is nonlinear. To solve the model, we use the nonlinear algorithm proposed by Jeffrey Fuhrer and Hoyt Bleakley (1996).

TABLE 2—PARAMETER CALIBRATION

	Parameter values
β	0.993
α_κ	0.360
δ	0.025
λ_w	1.150
ξ_p	0.350
η	0.650
σ_a	0.100
α_1	0.600
α_2	0.970
α_3	0.450
α_4	0.300
a_π	2.000
a_y	0.600
ρ	0.920

Table 2 summarizes the benchmark values used to calibrate the model. The parameters in the policy reaction function (12) are in line with the estimates reported in the literature (see, e.g., Fuhrer 2000). The interest-smoothing parameter ρ is set at 0.92. The coefficients on output and inflation are calibrated at $a_\pi = 2$ and $a_y = 0.6$. We follow Christiano et al. (2005) closely in the calibration of the remaining parameters. The discount factor β is set at $1.03^{-0.25}$, which implies a steady-state annualized real interest rate of 3 percent. The capital share α is set at 0.36. The depreciation rate δ is equal to 0.025, which implies an annual rate of depreciation on capital of 10 percent. As in Christiano et al. (2005), we set σ_a, the inverse of the elasticity of capital utilization with respect to the rental rate of capital, at 0.1. We maintain a similar degree of habit persistence in consumption, with the parameter describing the degree of habit persistence η set at 0.65; χ is set at 0.2, consistent with evidence reported by Fuhrer (2000). Investment adjustment costs are calibrated so that a permanent 1 percent decline in the price of capital induces a 20 percent change in investment. The parameter governing exogenous price rigidity, ξ_p, is equal to 0.35. This implies that price contracts last, on average, about 4.5 months, a value consistent with recent microeconomic evidence on the frequency of price changes (see Mark Bils and Peter Klenow 2004). Christiano et al. (2005) show that the model's response to a monetary policy shock is fairly insensitive to the value of this parameter; in particular, they show that the main results hold when prices are fully flexible ($\xi_p = 0$). This is because wage contracts, not price contracts, are the important nominal rigidity for imparting empirically consistent dynamics to the model.[33]

In calibrating the values for the probability of resetting wages in a given quarter, we follow an approach consistent with the parametrization in Christiano et al. (2005). In their setup, the probability of wage changes across quarters is identical and such that the average frequency of wage changes is approximately 2.5 quarters. Our way of calibrating the α_k's ensures that the average frequency of wage changes is about 2.1 quarters.[34] To capture the seasonal effects, we set the values of α_k so that $\alpha_4 < \alpha_3 < \alpha_1 < \alpha_2$. This corresponds to a situation in which a larger fraction of wages is changed during the course of the fourth and third quarters, a smaller fraction is changed in the first quarter, and an even smaller fraction during the second quarter. This assumption is in line with anecdotal evidence suggesting that wages are reset at the end or at the very beginning of the calendar year, with fewer changes taking place during the second quarter. The calibrated probabilities imply that 24 percent of the wage changes take place in the first quarter, 2 percent in the second quarter, 32 percent in the third quarter, and 42 percent in the fourth quarter.[35] These frequencies are consistent with wage-setting practices of New England firms surveyed in the Federal Reserve System's Beige Book and the additional sources of anecdotal evidence discussed in the introduction.[36]

E. *Model Results*

We now present the model-generated impulse responses to a 25-basis-point decline in the nominal interest rate on impact for output, inflation, the nominal interest rate, hours, and the real wage. To make the results comparable to the identifying assumption underlying our empirical exercise, we assume that the shock occurs at the end of period t, when all the period t nonpolicy variables have been already set. The model responses are plotted against the corresponding empirical responses and the 95 percent confidence bands from the previous section's estimated quarter-dependent VAR.

The responses of output are shown in Figure 14. Policy shocks occurring in the first and second quarters of the calendar year have a significant effect on output, since few households are allowed to reset their wages optimally. In the third and fourth quarters, the response of output is less than half the size of the response in the first half of the year. This difference in the response of output occurs even though our model features the same degree of real rigidities as in Christiano et al. (2005). In the model, real rigidities work in the direction of dampening the effect of uneven staggering of wages, but our calibrated values for the α_k's still imply relevant differences in the response of output across quarters. At the same time, real rigidities help to generate persistent responses to the policy shock. In line with Christiano et al. (2005), the model-generated impulse response for output peaks three to four quarters after the shock, depending on the quarter in which the shock occurs. This matches the response of output to a policy shock in the second quarter, although the output response to a policy shock in the first quarter is not as persistent as in the data.

The responses of inflation are displayed in Figure 15. Consistent with the data, the model produces impulse responses that are not strikingly different across quarters. In the two quarters after the shock, prices increase somewhat faster when the shock occurs in the third or fourth quarter—a feature we also observe in the data. The model responses fall comfortably

[33] David Card (1990) also provides empirical evidence on the importance of nominal wage rigidities in the transmission of aggregate demand shocks to real economic activity.

[34] In other words, $\alpha^4 \approx \prod_{k=1}^{4} \alpha_k$, where α is the constant probability value used in Christiano et al. (2005), equal to the geometric average of the quarter-dependent probabilities. Given our calibrated values for the α_k's in Table 2, this implies an average frequency of wage adjustment of 2.13 ($=1/(1 - \sqrt[4]{\prod_{k=1}^{4} \alpha_k}$) quarters.

[35] In each quarter k, the proportion of wage changes relative to the total number of changes in a given calendar year is $(1 - \alpha_k)/(4 - \sum_{k=1}^{4} \alpha_k)$.

[36] Findings about wage-setting practices of New England firms surveyed in the Federal Reserve System's Beige Book over the course of 2003 are available from the authors upon request. Firms are identified by sector, but the identity of the firms is kept confidential.

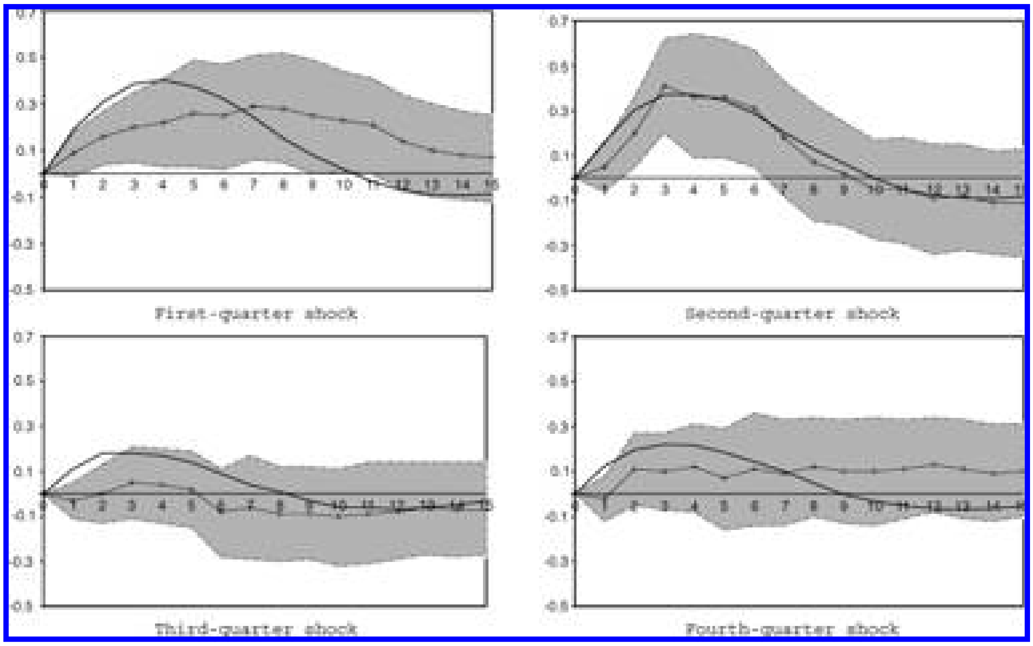

FIGURE 14. MODEL AND VAR IMPULSE RESPONSES OF OUTPUT TO A 25-BASIS-POINT DECLINE IN FED FUNDS RATE

Notes: Bold solid lines are the theoretical responses and solid lines-plus sign are the VAR responses. Broken lines indicate the 95 percent confidence intervals around VAR estimates. Vertical axis units are deviations from the steady-state path.

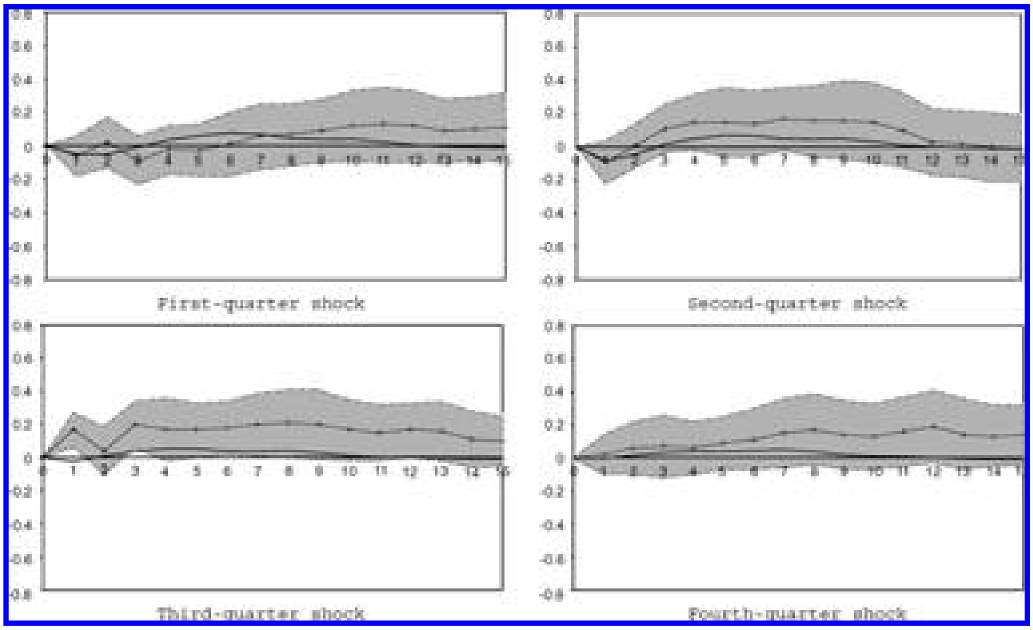

FIGURE 15. MODEL AND VAR IMPULSE RESPONSES OF INFLATION TO A 25-BASIS-POINT DECLINE IN FED FUNDS RATE

Notes: Bold solid lines are the theoretical responses and solid lines-plus sign are the VAR responses. Broken lines indicate the 95 percent confidence intervals around VAR estimates. Vertical axis units are deviations from the steady-state path (annualized percentage points).

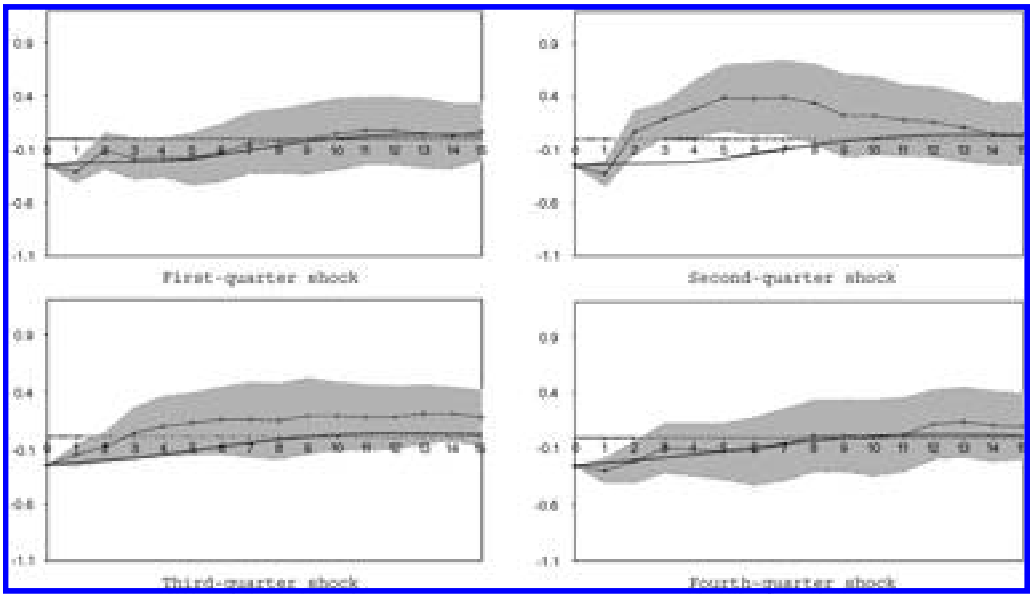

FIGURE 16. MODEL AND VAR IMPULSE RESPONSES OF THE FED FUNDS RATE TO A 25-BASIS-POINT DECLINE IN FED FUNDS RATE

Notes: Bold solid lines are the theoretical responses and solid lines-plus sign are the VAR responses. Broken lines indicate the 95 percent confidence intervals around VAR estimates. Vertical axis units are deviations from the steady-state path (annualized percentage points).

within the 95 percent confidence bands in all four quarters.

The responses of the federal funds rate, shown in Figure 16, are generally consistent with the empirical responses, with the exception of the response to a second-quarter shock, when the theoretical response is weaker than its empirical counterpart.

As for the responses of the real wage, the model's setup with Calvo-style wage setting implies that at each point in time the aggregate wage is a weighted average of the aggregate wage prevailing in the previous period (the "preset" wage) and of the "optimal" wage, that is, the wage set by workers who at that particular point in time are able to adjust their wage optimally. Figure 17 shows the model responses of the aggregate and the optimal real wage. There are noticeable differences in the response of the optimal wage across quarters. Still, consistent with our empirical findings, the responses of the real aggregate wage are fairly similar across quarters. The intuition for this result is the following. After an expansionary monetary shock, aggregate demand goes up and so does the derived demand for labor. In a period when a large fraction of workers cannot readjust their wage optimally, these workers will have to supply as much labor as firms demand at the preset wage. The fewer workers who can adjust the wage optimally will raise the wage significantly following the shock.[37] Specifically, they will take into account that (a) activity is responding strongly to the shock, with firms demanding more labor from the many workers having a preset wage; and (b) in future periods many other workers will reset their wages at a higher level, given the persistent effect of the shock on aggregate demand. The aggregate wage is then determined by a large fraction of preset wages and a small fraction of reoptimized wages responding strongly to the shock.

In contrast, in a period in which a large fraction of the workers can adjust their wage optimally, firms can no longer count on as much

[37] Recall that due to the Dixit-Stiglitz formulation, different types of workers are not perfect substitutes. Given the higher levels of aggregate demand, the demand for the workers who can reset wages also goes up, and hence their wages.

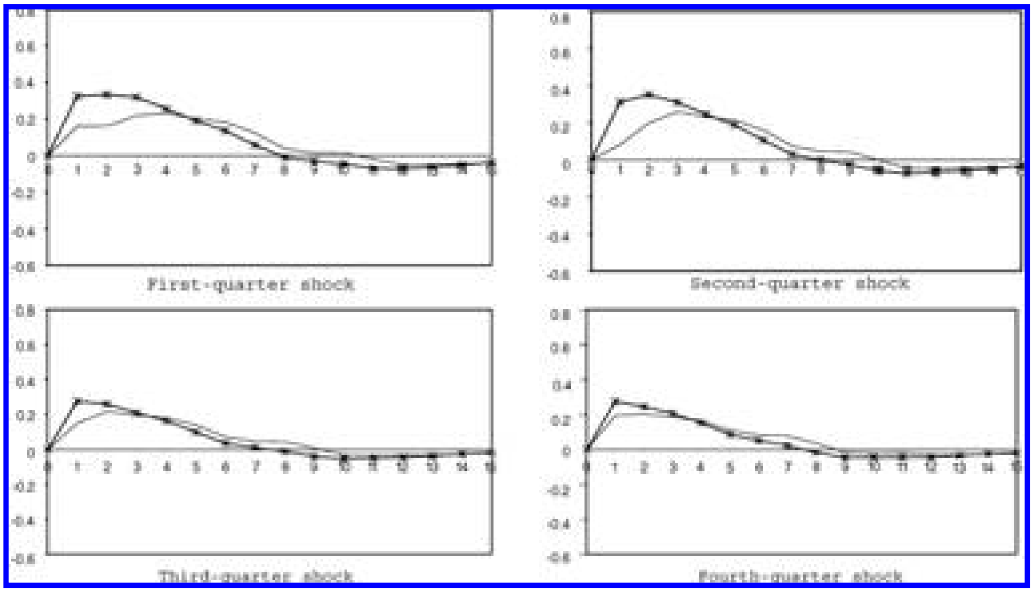

FIGURE 17. MODEL RESPONSES OF THE AGGREGATE REAL WAGE (w_t) AND THE OPTIMAL REAL WAGE ($\hat{\hat{w}}_t$) TO A 25-BASIS-POINT DECLINE IN FED FUNDS RATE

Notes: Solid lines are the responses of the aggregate real wage and solid lines-plus asterisks are the responses of the optimal real wage.

labor at preset wages as in the previous scenario. Now firms meet an upward-sloping supply of labor for a large number of workers. The increase in aggregate activity following an expansionary monetary policy shock is therefore smaller, and the economy operates at a level closer to the flexible-wage-price potential. Because aggregate activity stays closer to its long-run equilibrium following the shock, the optimal wage does not increase by as much as in the previous case. The aggregate real wage is now determined by a small fraction of preset wages and a large fraction of reoptimized wages, with the latter responding less strongly to the shock than in the previous scenario.

In sum, the model predicts a quarter-dependent response for the optimal wage, but the response of the aggregate wage needs not be strikingly different across different quarters, precisely because when the optimal wage responds more (less), the optimal wage has a smaller (larger) weight in the aggregate wage. In Figure 18 we plot the model-generated responses for the aggregate real wage, together with the corresponding empirical responses. As the figure shows, the theoretical responses fall consistently within the empirical confidence bands.[38]

The adjustment of labor services plays a central role in the transmission mechanism in the presence of nominal wage rigidities. After an expansionary monetary policy shock, firms use labor inputs more intensively, increasing the number of hours. The larger the extent of wage rigidity, the greater the intensity with which labor inputs are used. The model thus predicts that in quarters when nominal wages are relatively rigid (the first two quarters of the calendar year), labor hours will respond to a monetary policy shock more strongly than in quarters when nominal wages are relatively flexible (the third and fourth quarters of the calendar year). In Figure 19, we plot both the theoretical and empirical impulse responses of hours to a monetary expansion. In line with the data, the the-

[38] Note that as in Christiano et al. (2005), real wages are procyclical. This is because prices, which are a markup over marginal costs, are a weighted average of wages and the interest rate R. In an expansion, R declines and hence the price level increases less than wages.

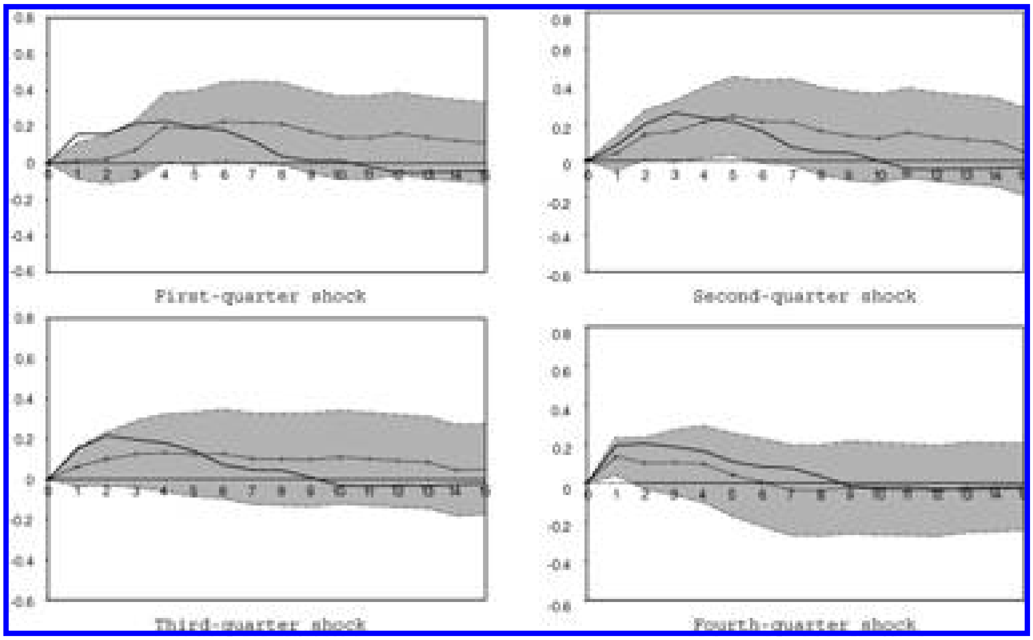

FIGURE 18. MODEL AND VAR IMPULSE RESPONSES OF THE AGGREGATE REAL WAGE TO A 25-BASIS-POINT DECLINE IN FED FUNDS RATE

Notes: Bold solid lines are the theoretical responses and solid lines-plus sign are the VAR responses. Broken lines indicate the 95 percent confidence intervals around VAR estimates. Vertical axis units are deviations from the steady-state path.

oretical response is strongly seasonal, being on average twice as large when the shock takes place in the first two quarters. Though the theoretical responses tend to be slightly smaller than the empirical ones, they usually fall within the empirical confidence bands.

In sum, the model results indicate that a realistic amount of uneven staggering yields to different output and hours dynamics, depending on the quarter in which the policy shock takes place. The differences occur even if the model embeds a significant (and empirically plausible) degree of real rigidities. The responses for several key model variables are not only qualitatively, but also quantitatively, consistent with the quarter-dependent patterns we have documented in actual data.

IV. Concluding Remarks

The paper documents novel findings regarding the impact of monetary policy shocks on real activity. After a monetary expansion that takes place in the first quarter of the year, output picks up quickly and tends to die out at a relatively fast pace. This pattern is even more pronounced when the monetary policy expansion takes place in the second quarter of the year. In contrast, output responds little when the monetary expansion takes place in the third and fourth quarters of the year. The conventional finding that monetary shocks affect output with long delays, and that the effect is highly persistent, may be interpreted as the combination of these different output impulse responses.

We argue that the differential responses are not driven by different types of monetary policy interventions, nor by different "states" of the economy across quarters. Encouraged by anecdotal evidence on the timing of wage changes, which suggests that a large fraction of wages are reset toward the end of each calendar year, we propose a potential explanation for the differential responses based on contractual lumping and develop a theoretical general equilibrium model based on Christiano et al. (2005) featuring uneven staggering of nominal wage contracts. The model generates impulse responses that quantitatively match those found in the data.

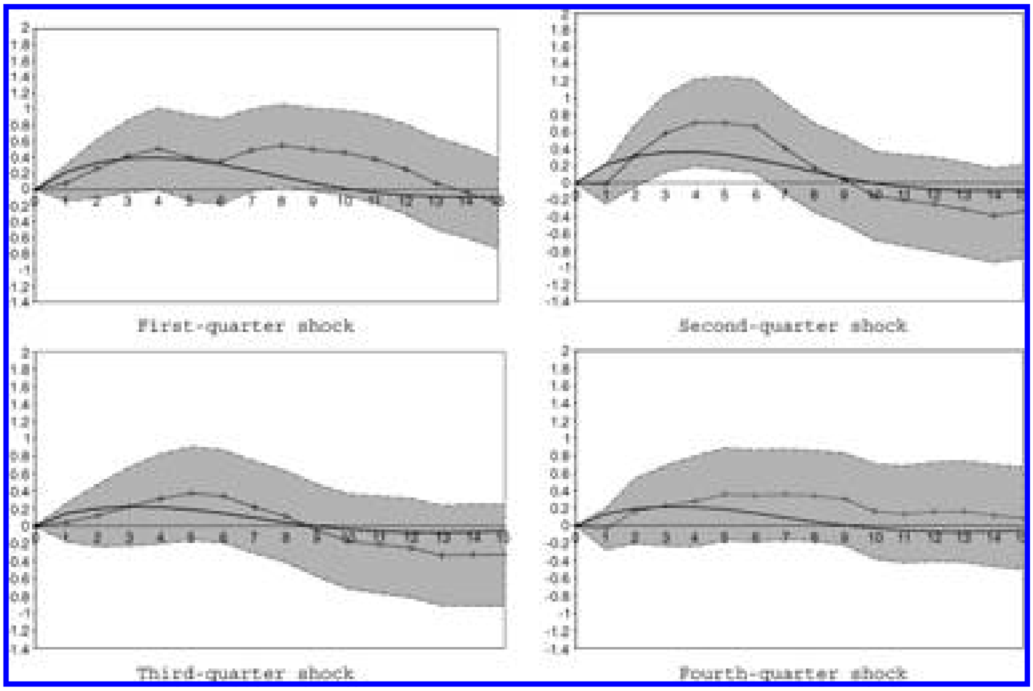

FIGURE 19. MODEL AND VAR IMPULSE RESPONSES OF HOURS TO A 25-BASIS-POINT DECLINE IN FED FUNDS RATE

Notes: Bold solid lines are the theoretical responses and solid lines-plus sign are the VAR responses. Broken lines indicate the 95 percent confidence intervals around VAR estimates. Vertical axis units are deviations from the steady-state path.

While our model assumes uneven staggering, there are studies in the literature addressing the optimality of uniform staggering versus synchronization of price (or wage) changes. The general finding of this literature is that synchronization is the equilibrium timing in many simple Keynesian models of the business cycle.[39] Yet, the new generation of Keynesian models has glossed over this finding and assumed uniform staggering as both a convenient modeling tool and an essential element in the transmission mechanism of monetary policy shocks. This paper provides an empirical setting to test the hypothesis of uniform versus uneven staggering of wage changes, and in so doing argues for the empirical and theoretical relevance of models in which wage changes are less staggered and more synchronized.

We have addressed the robustness of our findings along several dimensions, but additional evidence could corroborate our results. Other shocks, such as technology shocks, could also have a different impact across the calendar year if wage staggering is not uniform. However, while there is some consensus on how to identify monetary policy shocks, the identification of other types of shocks remains contentious and would require additional variables in our VAR, thus reducing degrees of freedom at the estimation stage. A more promising avenue, in our view, is the examination of international evidence. Other countries likely exhibit uneven staggering of wage contracts, and the transmission mechanism of monetary shocks should be affected accordingly. In this respect, Japan is particularly relevant, because a large fraction of Japanese firms set wages in the

[39] Lawrence Ball and Stephen Cecchetti (1988) show that staggering can be the equilibrium outcome in some settings with imperfect information, but even then such a result is not necessarily pervasive, since it depends on the structure of the market in which firms compete and on firms setting prices for a very short period of time. In other settings, staggering can be the optimal outcome for wage negotiations if the number of firms is very small (see Gary Fethke and Andrew Policano 1986). The incentive for firms to stagger wage negotiation dates, however, diminishes the larger the number of firms in an economy.

spring season (during the wage-setting process known as Shunto, or "spring offensive"). Preliminary findings for Japan indicate that monetary policy shocks that take place in the third quarter (after the Shunto) have a large impact on output, whereas monetary policy shocks occurring in the second quarter (during the Shunto) have virtually no output effect.

Appendix

This Appendix is composed of two parts. The first part details the derivation of the equations governing the wage dynamics (equations (10) and (11) in the text). The second part presents the model's log-linearized relationships for the nonpolicy variables which, together with the policy equation (12) in the text, summarize the model's dynamics around its steady state.

Wage-Setting Process

We let period t correspond to the first quarter (and so does $t + 4j$, $\forall j$) while $t + 1 + 4j$, $\forall j$ corresponds to the second, $t + 2 + 4j$, $\forall j$ corresponds to the third, and $t + 3 + 4j$, $\forall j$ corresponds to the fourth quarter. Since all labor types resetting their wages in a given quarter will choose the same wage, we can drop the i's from the notation and simply refer to the optimal wage in period t as $\tilde{W}_t$. The first-order condition for the optimal wage $\tilde{W}_t$ in the first quarter satisfies

$$E_{t-1}\left\{\begin{aligned}&\sum_{j=0}^{\infty}\prod_{i=1}^{4}\alpha_i^j\beta^{4j}h_{t+4j}\psi_{t+4j}\left\{\left[\frac{\tilde{W}_t\cdot X_{t,4j}}{P_{t+4j}}-\lambda_w\frac{z_{h,t+4j}}{\psi_{t+4j}}\right]\right.\\&+\alpha_2\beta h_{t+4j+1}\psi_{t+4j+1}\left[\frac{\tilde{W}_t\cdot X_{t,4j+1}}{P_{t+4j+1}}-\lambda_w\frac{z_{h,t+4j+1}}{\psi_{t+4j+1}}\right]\\&+\alpha_2\alpha_3\beta^2 h_{t+4j+2}\psi_{t+4j+2}\left[\frac{\tilde{W}_t\cdot X_{t,4j+2}}{P_{t+4j+2}}-\lambda_w\frac{z_{h,t+4j+2}}{\psi_{t+4j+2}}\right]\\&\left.+\alpha_2\alpha_3\alpha_4\beta^3 h_{t+4j+3}\psi_{t+4j+3}\left[\frac{\tilde{W}_t\cdot X_{t,4j+3}}{P_{t+4j+3}}-\lambda_w\frac{z_{h,t+4j+3}}{\psi_{t+4j+3}}\right]\right\}\end{aligned}\right\}=0,$$

where $z_{h,t+l}$ is the derivative of z with respect to h_{t+l} and $X_{t,l}$ is defined in (9). Log-linearizing this expression around its steady state, we obtain

$$\begin{aligned}\widehat{\tilde{W}}_t = \hat{P}_{t-1} + \Theta E_{t-1}\Bigg\{\sum_{j=0}^{\infty}\prod_{i=1}^{4}\alpha_i^j\beta^{4j}\cdot\Bigg\{\Bigg[&\hat{\pi}_{t+4j}+\frac{\lambda_w-1}{2\lambda_w-1}\left(\frac{\lambda_w}{\lambda_w-1}(\hat{W}_{t+4j}-\hat{P}_{t+4j})+\hat{L}_{t+4j}-\hat{\psi}_{t+4j}\right)\Bigg]\\&+\alpha_2\beta\left[\hat{\pi}_{t+4j+1}+\frac{\lambda_w-1}{2\lambda_w-1}\left(\frac{\lambda_w}{\lambda_w-1}(\hat{W}_{t+4j+1}-\hat{P}_{t+4j+1})+\hat{L}_{t+4j+1}-\hat{\psi}_{t+4j+1}\right)\right]\\&+\alpha_2\alpha_3\beta^2\left[\hat{\pi}_{t+4j+2}+\frac{\lambda_w-1}{2\lambda_w-1}\left(\frac{\lambda_w}{\lambda_w-1}(\hat{W}_{t+4j+2}-\hat{P}_{t+4j+2})+\hat{L}_{t+4j+2}-\hat{\psi}_{t+4j+2}\right)\right]\\&+\alpha_2\alpha_3\alpha_4\beta^3\left[\hat{\pi}_{t+4j+3}+\frac{\lambda_w-1}{2\lambda_w-1}\left(\frac{\lambda_w}{\lambda_w-1}(\hat{W}_{t+4j+3}-\hat{P}_{t+4j+3})+\hat{L}_{t+4j+3}-\hat{\psi}_{t+4j+3}\right)\right]\Bigg]\Bigg\}\Bigg\},\end{aligned}$$

where $\hat{x}_t = (x_t - x)/x$ and x is the nonstochastic steady-state level of x_t and

$$\Theta^{-1} = \sum_{j=0}^{\infty} \prod_{i=1}^{4} \alpha_i^j \beta^{4j} (1 + \alpha_2 \beta + \alpha_2 \alpha_3 \beta^2 + \alpha_2 \alpha_3 \alpha_4 \beta^3) = \frac{(1 + \alpha_2 \beta + \alpha_2 \alpha_3 \beta^2 + \alpha_2 \alpha_3 \alpha_4 \beta^3)}{1 - \alpha_1 \alpha_2 \alpha_3 \alpha_4 \beta^4}.$$

We can then write $\hat{\tilde{W}}_t$ in recursive form as

$$(13) \qquad \hat{\tilde{W}}_t = \hat{P}_{t-1} + \Theta\left[\hat{\pi}_t + \frac{\lambda_w - 1}{2\lambda_w - 1}\left(\frac{\lambda_w}{\lambda_w - 1}(\hat{W}_t - \hat{P}_t) + \hat{L}_t - \hat{\psi}_t\right)\right]$$
$$+ \beta\alpha_2 \frac{1 + \beta\alpha_3(1 + \beta\alpha_4 + \beta^2\alpha_4\alpha_1)}{1 + \beta\alpha_2(1 + \beta\alpha_3 + \beta^2\alpha_3\alpha_4)} E_{t-1}(\hat{\tilde{W}}_{t+1} - \hat{P}_t).$$

The aggregate wage level is a weighted average of the optimal wage set by the workers who received the signal to reoptimize, and the wage of the workers who did not get the signal. As mentioned in the text, the workers who do not reoptimize have their wage indexed to the previous period rate of inflation. If we keep our convention that t (and $\{t + 4j\}$) corresponds to the first quarter, the proportion of workers changing wages at t will be equal to $(1 - \alpha_1)$. Then, the expression for the aggregate wage level is given by

$$(14) \qquad \hat{W}_t = (1 - \alpha_1)\hat{\tilde{W}}_t + \alpha_1(\hat{W}_{t-1} + \hat{\pi}_{t-1}).$$

Equations (13) and (14) describe the law of motion for the optimal wage and the aggregate wage in the first quarter of the calendar year. To generalize these expressions to any quarter of the calendar year, we introduce the dummy variables q_{kt}, with $k = 1, \cdots, 4$ which take on the value one in the kth quarter and zero otherwise. We can then write the equations governing the wage dynamics as

$$(15) \qquad \hat{\tilde{W}}_t = \hat{P}_{t-1} + \frac{1 - \prod_{j=1}^{4} \alpha_j \beta^4}{\Gamma_{(q,\alpha,\beta)}}\left[\hat{\pi}_t + \frac{\lambda_w - 1}{2\lambda_w - 1}\left(\frac{\lambda_w}{\lambda_w - 1}(\hat{W}_t - \hat{P}_t) + \hat{L}_t - \hat{\psi}_t\right)\right]$$
$$+ \beta \sum_{k=1}^{4} q_{kt}\alpha_{k+1} \frac{\Phi_{(q,\alpha,\beta)}}{\Gamma_{(q,\alpha,\beta)}} E_{t-1}(\hat{\tilde{W}}_{t+1} - \hat{P}_t),$$

and

$$(16) \qquad \hat{W}_t = \left[1 - \left(\sum_{k=1}^{4} q_{kt}\alpha_k\right)\right]\hat{\tilde{W}} + \left(\sum_{k=1}^{4} q_{kt}\alpha_k\right)(\hat{W}_{t-1} + \hat{\pi}_{t-1}),$$

where

$$\Gamma_{(q,\alpha,\beta)} = 1 + \beta\{q_{1t}\alpha_2[1 + \beta\alpha_3 + \beta^2\alpha_3\alpha_4] + q_{2t}\alpha_3[1 + \beta\alpha_4 + \beta^2\alpha_4\alpha_1]$$
$$+ q_{3t}\alpha_4[1 + \beta\alpha_1 + \beta^2\alpha_1\alpha_2] + q_{4t}\alpha_1[1 + \beta\alpha_2 + \beta^2\alpha_2\alpha_3]\};$$

$$\Phi_{(q,\alpha,\beta)} = 1 + \{\beta q_{1t}\alpha_3[1 + \beta\alpha_4 + \beta^2\alpha_4\alpha_1] + q_{2t}\alpha_4[1 + \beta\alpha_1 + \beta^2\alpha_1\alpha_2]$$
$$+ q_{3t}\alpha_1[1 + \beta\alpha_4 + \beta^2\alpha_4\alpha_3] + q_{4t}\alpha_2[1 + \beta\alpha_3 + \beta^2\alpha_3\alpha_4]\}.$$

Using the definitions of $\hat{\hat{w}}_t$ and $\hat{w}_t$ in the text ($\hat{\hat{w}}_t \equiv \hat{\hat{W}}_t - \hat{P}_t$ and $\hat{w}_t \equiv \hat{W}_t - \hat{P}_t$), equations (15) and (16) become, respectively,

$$\hat{\hat{w}}_t + \hat{\pi}_t = \frac{1 - \prod_{j=1}^{4} \alpha_j \beta^4}{\Gamma_{(q,\alpha,\beta)}} \left[\hat{\pi}_t + \frac{\lambda_w - 1}{2\lambda_w - 1} \left(\frac{\lambda_w}{\lambda_w - 1} \hat{w}_t + \hat{L}_t - \hat{\psi}_t \right) \right]$$

$$+ \beta \sum_{k=1}^{4} q_{kt} \alpha_{k+1} \frac{\Phi_{(q,\alpha,\beta)}}{\Gamma_{(q,\alpha,\beta)}} E_{t-1} (\hat{\hat{w}}_{t+1} + \hat{\pi}_{t+1}),$$

and

$$\hat{w}_t = \left[1 - \left(\sum_{k=1}^{4} q_{kt} \alpha_k \right) \right] \hat{\hat{w}}_t + \left(\sum_{k=1}^{4} q_{kt} \alpha_k \right) (\hat{w}_{t-1} + \hat{\pi}_{t-1} - \hat{\pi}_t),$$

which correspond to equations (10) and (11) in the text.

The Model's Log-Linearized Equilibrium Equations

We now briefly describe the set of log-linearized equations which, together with the wage equations (10) and (11), and the Taylor rule (12), characterize the dynamics of the model.

The inflation dynamics implied by Calvo-pricing takes the form

$$\text{(17)} \qquad \hat{\pi}_t = \frac{1}{1+\beta} \hat{\pi}_{t-1} + E_{t-1} \left\{ \frac{\beta}{1+\beta} \hat{\pi}_{t+1} + \frac{(1 - \beta\xi_p)(1 - \xi_p)}{(1+\beta)\xi_p} \hat{s}_t \right\},$$

where the marginal cost, $\hat{s}_t$, is given by

$$\text{(18)} \qquad \hat{s}_t = \alpha(\hat{w}_t + \hat{R}_t + \hat{L}_t - \hat{k}_t) + (1 - \alpha)(\hat{w}_t + \hat{R}_t).$$

Firms that do not reoptimize in a given period change prices according to the most recent rate of inflation, hence the presence of lagged inflation in the equation. Because in the Calvo setup firms cannot adjust prices optimally every period, inflation is a function not only of current marginal costs, but also of the expected present discounted value of current and future marginal costs (entering through $\hat{\pi}_{t+1}$).

The Euler equation for consumption is given by

$$\text{(19)} \qquad E_{t-1} \left\{ \beta\chi\hat{\psi}_{t+1} + \sigma_c \left(\hat{c}_t - \frac{b}{1-\chi} \hat{e}_t \right) + (b + \chi)\beta\sigma_c \left(\hat{c}_{t+1} - \frac{b}{1-\chi} \hat{e}_{t+1} \right) \right\} = \hat{\psi}_t,$$

where $\sigma_c = (1 - \chi/1 - \chi - b)(1 - \beta\chi/1 - \beta\chi - \beta b)$, $\hat{e}_{t+1} = \chi\hat{e}_t + (1 - \chi)\hat{c}_t$, and

$$\text{(20)} \qquad E_{t-1}\hat{\psi}_{t+1} = \hat{\psi}_t - E_{t-1}(\hat{R}_{t+1} - \hat{\pi}_{t+1}).$$

The presence of habit formation implies that consumers' current utility is determined by current consumption relative to a reference level of consumption. The parameter χ indexes the persistence or "memory" in the habit-formation reference level. In essence, the Euler equation says that the expected level of consumption next period relative to its reference level depends on current consumption relative to its reference level on a present-discounted stream of expected real interest rates.

The specification for investment takes the form

$$E_{t-1}(\hat{\iota}_t - \hat{\iota}_{t-1}) = \beta E_{t-1}(\hat{\iota}_{t+1} - \hat{\iota}_t) + \frac{1}{\kappa}\hat{P}^k_t, \tag{21}$$

where $\hat{P}^k_t$ is the hypothetical spot price of a unit of capital stock installed at time t. In turn, the price of capital is determined according to a familiar asset-pricing equation whereby the price of capital at time t depends on the price of capital next period plus its period t dividend, equal to the rental rate of capital $\hat{r}^k_t = \hat{w}_t + \hat{R}_t + \hat{L}_t - \hat{k}_t$. Forward iteration of the equation for the price of capital yields the following relationship:

$$E_{t-1}\hat{P}^k_t = (1 - \beta(1-\delta))E_{t-1}\sum_{i=0}^{\infty}\beta^i(1-\delta)^i\hat{r}^k_{t+1+i} - E_{t-1}\sum_{i=0}^{\infty}\beta^i(1-\delta)^i(\hat{R}_{t+1+i} - \hat{\pi}_{t+1+i}). \tag{22}$$

This relationship can be substituted into (21) to obtain an expression for the change in investment in terms of a present discounted value of the rental rate of capital and of the real interest rate.

The Euler equation for the household's capital-utilization decision becomes

$$E_{t-1}u_t = E_{t-1}(\hat{k}_t - \hat{\bar{k}}_t) = \frac{1}{\sigma_a}\hat{r}^k_t. \tag{23}$$

This expression says that the expected marginal benefit of raising utilization must equal the associated expected marginal cost.

Finally, the log-linearized versions of the capital accumulation equation and the economy-wide resource constraint are given by

$$\hat{\bar{k}}_t = (1-\delta)\hat{\bar{k}}_{t-1} + \delta\hat{\iota}_t, \tag{24}$$

and

$$s_c\hat{c}_t + s_g\hat{g}_t + s_i\left\{\hat{\iota}_t + \left[\frac{1}{\beta} - (1-\delta)\right]\frac{1}{\delta}(\hat{k}_t - \hat{\bar{k}}_t)\right\} = \hat{y}_t, \tag{25}$$

where output $\hat{y}_t$ is equal to

$$\hat{y}_t = \alpha\hat{k}_t + (1-\alpha)\hat{L}_t. \tag{26}$$

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