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The Timing of Monetary Policy Shocks

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Abstract:
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 stylized dynamic general equilibrium model, we show that a very modest amount of uneven staggering can generate differences in output responses similar to those found in the data.

JEL Codes: E1, E52, E58, E32, E31

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1 Introduction

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, Christiano, Eichenbaum, and 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 (Taylor 1980 and 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

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

technology sector reveal that more than 90 percent of the companies use a focal base-pay administration with annual pay-change reviews; pay changes usually take place at the beginning of the new fiscal year. According to the same survey, 60 percent of the IT 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 Russell 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 quarter.[3] Finally, reports on collective bargaining activity compiled by the Bureau of Labor Statistics indicate that the distribution of expirations and wage reopening dates is tilted towards the second semester of the year.[4,5]

If the staggering of wage contracts is not uniform, as the anecdotal evidence seems to suggest, 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 structural 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 such a 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

[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 & Poor's COMPUSTAT. To compute the percentages, we weighted each firm in the Russell 3,000 index by the firm's number of employees.

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

[5]As concerns pricing practices, Zbaracki Ritson, Levy, Dutta, and Bergen (2004) present evidence that prices in the manufacturing sector are changed once a year. Typically, new prices take effect at the turn of the calendar year.

shock. The response is even more front-loaded and dies out faster when the shock takes place in the second quarter. Then, the peak effect is reached three quarters after the shock. In both the first and second quarters of the calendar year, 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 eight quarters by an average of about one quarter of one percent. 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-point unexpected monetary expansion in either the third or fourth quarter raises output in the eight quarters following the shock by less than one tenth of one percent 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 can then 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 wages. The price and the wage responses are delayed when the shock occurs in the first half of the year, whereas prices and wages respond more quickly when the shock occurs in the second half of the year.

We interpret the differential responses across quarters in the context of a simple stochastic dynamic general equilibrium model that allows for uneven staggering of wage contracts. To represent uneven staggering, the model features a variant of the mechanism proposed by Calvo (1983). The crucial difference is that the probability of changing wages is not constant across quarters. We show that a modest amount of uneven staggering can lead to significantly different output responses. This happens even if the cumulative effect of the monetary policy shock on wages and prices is not strikingly different across quarters, as appears to be the case empirically.

The remainder of the paper is organized as follows. Section 2 presents the empirical methodology and introduces the data. Section 3 presents the dynamic effects of monetary policy on different macroeconomic aggregates and performs a set of robustness tests. Section 4 illustrates the theoretical model and discusses its implications in light of the empirical findings. Section 5 offers some concluding remarks.

2 Methodology

2.1 Empirical Model

Our empirical analysis for measuring the effect of monetary policy shocks relies on a very general structural model of the macroeconomy represented by the following system of equations:

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

$$p_t = \sum_{s=0}^{k} \mathbf{D}_s \mathbf{Y}_{t-s} + \sum_{s=1}^{k} \mathbf{G}_s \; p_{t-s} + \mathbf{A}^p v_t^p. \tag{2}$$

Boldface letters indicate vectors or matrices of variables or coefficients. Specifically, $\mathbf{Y}_t$ is a vector of non-policy macroeconomic variables (e.g., output, prices, and wages), and p_t is the 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. Equation (1) allows the non-policy 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, $\mathbf{v}^y$. 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.[6,7] As such, the system embeds the key assumption 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 many recent VAR analyses.[8]

The structural 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 non-policy 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 non-policy variables, $\mathbf{Y}_t$.[9]

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:

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

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

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

[7]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).

[8]See, among others, Bernanke and Blinder (1992), Rotemberg and Woodford (1997), Bernanke and Mihov (1998), Christiano, Eichenbaum, and Evans (1999), and Boivin and Giannoni (2003).

[9]Note also that 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 $\mathbf{Y}_t$ already mentioned. We are unaware of any evidence suggesting that policy responses to given outcomes vary by season.

by-equation using ordinary least squares. The effect of policy innovations on the non-policy 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. An estimated series for the policy shock can be obtained via a Choleski decomposition of the covariance matrix of the reduced-form residuals.

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

$$\Xi_0 \mathbf{X}(\mathbf{t}) = \Xi_1 \mathbf{X}(\mathbf{t-1}) + \mathbf{U}(\mathbf{t}), \tag{4}$$

where Ξ_0 and Ξ_1 are parameter matrices containing the parameters in $\mathbf{A}(L, q)$ in (3), and $\mathbf{U}(\mathbf{t}) = (\mathbf{U}_{1,t}, \mathbf{U}_{2,t}, \mathbf{U}_{3,t}, \mathbf{U}_{4,t})$, 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, the matrix Ξ_0 is lower block-triangular and can be inverted to yield:

$$\mathbf{X}(\mathbf{t}) = \Xi_0^{-1} \Xi_1 \mathbf{X}(\mathbf{t-1}) + \Xi_0^{-1} \mathbf{U}(\mathbf{t}). \tag{5}$$

The inverse of a lower block-triangular matrix is still a lower block-triangular matrix. The system (5) then illustrates that when a monetary policy shock occurs in the first quarter, the response of the non-policy variables in the next quarter will be governed by the reduced-form dynamics of the non-policy variables in the second quarter. The response two quarters after the initial shock will be

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

governed by the reduced-form dynamics of the non-policy variables in the third quarter, and so on.

2.2 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. However, even if F-tests reject the null hypothesis of no time dependence, this does not ensure that the impulse responses generated by the quarter-dependent VAR are statistically different from the responses generated by the standard VAR. Impulse-response functions are nonlinear combinations of the estimated coefficients in the VAR, and as a result F-tests on the linear reduced-form VAR do not map one-for-one into a test on the impulse-responses.

For this reason, we assess the significance of quarter-dependence on the impulse-response functions directly. 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. That is:

$$D = \sup |x_t^q - x_t|$$

where x_t^q denotes the period t response in the quarter-dependent model and x_t the response in the standard, non-time-dependent model.[11] Under the assumption that the structural monetary shock is normally distributed, impulse responses are asymptotically normal, and so is the D-statistic. We

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

construct an empirical distribution of D by bootstrapping the residuals of the reduced-form, non-time-dependent VAR. At each draw, we generate a new data set and estimate new impulse responses from both the quarter-dependent and the standard VAR. This yields a new value for 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.

2.3 Data and Estimation

Our benchmark analysis is based on seasonally-adjusted 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. The non-policy variables in the system include real GDP, the GDP deflator, and an index of spot commodity prices.[12] 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 the robustness section we replace the GDP deflator by the core CPI and by an index of wages, given by compensation per-hour in the nonfarm business sector.[13] The robustness section also presents results based on non-seasonally adjusted data.

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 non-policy variables as deterministic, and allow for a linear trend in each of the equations of the

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

[13]The source for both the core CPI and compensation per-hour in the nonfarm business sector is the Bureau of Labor Statistics.

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.

3 The Dynamic Effects of Monetary Policy Shocks

3.1 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, with finer lines denoting the 80 percent confidence band around the estimated responses.[14] 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 responses of the variables to the shock are 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 of the peak response twelve quarters after the shock. The center panel shows that prices start to rise reliably three quarters after the shock, although it takes about one year and a half 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

[14] Much applied work uses 95 percent confidence intervals. Sims and Zha (1999) note that the use of high-probability intervals camouflages the occurence 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 1.3 standard error in the Gaussian case.

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: twelve quarters after the shock, the response of output is less than a third of the peak response, which occurs seven 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 fed 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 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 funds rate in fact moves into positive territory three quarters after the shock, and stays significantly positive for an additional six quarters.

The responses to a monetary policy shock in either the third or the fourth quarter of the year stand in sharp contrast to 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.

It is possible to assess directly from the confidence intervals in Figures 2 to 5 that the output responses to a policy shock in the first versus the the second half of the calendar year are statistically different. Specifically, the estimated output response in the first and second quarters lie outside the confidence interval of the output response in the third quarter. The estimated output response in the third quarter in turn falls outside the confidence interval of the estimated responses in the first and second quarters. In addition, the output response in the second quarter lies outside the confidence interval of the output response in the fourth quarter (and vice-versa) for six quarters after the shock. Such differences are significant from an economic standpoint, too. The policy shock raises output in the following eight quarters by an average of about one quarter of one percent in either the first or the second quarter. In contrast, the increase in output is less than one tenth of one percent on average in both the third and the fourth quarter.

The difference in impulse responses documented in figures 1 to 5 is corroborated by other tests on the importance of quarter-dependence. Equation-by-equation F-tests in the reduced-form VAR (3) yield p-values of 0.20 for the output equation, 0.06 for the price equation, 0.03 the for the commodity prices equation, and 0.004 for the federal funds rate equation. While suggestive of the existence of time dependence, these relatively low p-values do not necessarily translate into statistically different impulse responses. For this purpose, we evaluate the D-statistic described in Section 2.2, which assesses whether the maximum difference between each impulse response in the quarter-dependent VAR and the corresponding response 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 by this statistic the output response in the first and second 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 6 percent level in the third quarter, while in the fourth quarter the null hypothesis cannot be rejected at standard confidence levels. As it is also possible to infer from Figures 1 through 5, the table shows that there is little evidence in favor of quarter-dependent price impulse responses except in the third quarter, when prices start to rise significantly immediately after the shock (center panel of Figure 4). In the third quarter, the null hypothesis of a price response equal to the non-time-dependent response is rejected at the asymptotic 10 percent level.

3.2 The Distribution of Monetary Policy Shocks and the State of the Economy

An important issue to consider is whether the different impulse responses we obtain across quarters are the result of different types of shocks. In principle, differences in the intensity and direction (expansionary versus contractionary) of shocks could result in different impulse responses. To explore this hypothesis, 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 0.27; the largest p-value corresponds to the test between the second and third quarters and is equal to 0.96. These findings suggest that differences in the type of monetary policy shocks across quarters are unlikely to provide an explanation for the quarterly differences in impulse responses documented in Figures 2 to 5.

Another 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.[15] 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. Barsky and 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 (with 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 in the first (recession) and second (expansion) quarters, and the response is weak in the third (recession) and fourth (expansion) quarters. As we show in the robustness section, such a pattern continues to hold when we use seasonally unadjusted data. We thus conclude that seasonal fluctuations in output are unlikely to drive our findings.

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[15]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.

3.3 Robustness

We now discuss some exercises pertaining to the robustness of our findings. Since the proposed explanation for the different impulse-responses relies on uneven staggering of wage contracts across quarters, we replaced the GDP deflator with a wage index in the vector of non-policy variables, $\mathbf{Y}$. Using wages in lieu of final prices does not alter our main findings. Figures 6 to 9 show the responses of output, wages, and the federal funds rate to a 25-basis-point decline in the funds rate for the quarter-dependent reduced-form VAR (3). The estimated output responses 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 some interest to check whether our results are driven entirely by the seasonal adjustment. To this end, we estimated impulse-responses to a monetary shock from the quarter-dependent, reduced-form VAR (3) using seasonally unadjusted data for the non-policy variables, $\mathbf{Y}$.[16] The results from this exercise are illustrated in Figures 10 through 13, 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.

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,[17] and prices are expressed in log first differences. Such a

[16]Since there are no data on the seasonally unadjusted GDP deflator, we replace the GDP deflator with the seasonally unadjusted CPI.

[17]Specifically, we consider the deviation of log real GDP from its segmented deterministic linear trend, with break-

specification is common in the VAR literature (see, e.g., Boivin and Giannoni 2003). Figures 14 through 17 illustrate that the estimated impulse responses to a 25-basis-point decline in the funds rate 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.

3.4 Interpretation of the Evidence

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 slightly 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.

The rejection of the hypotheses that monetary policy shocks are different across quarters and that the state of the economy is triggering the differences in the impulse responses, together with the reported VAR results, lends itself to the following interpretation in the context of models with contractual rigidities. If a large number of firms sign wage contracts (or set prices) 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 prices. In contrast, monetary policy shocks in the second half of the year will be quickly followed by 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 show next that this intuition can be formalized in a dynamic, stochastic,

points in 1974 and 1995.

general equilibrium setup that allows for uneven staggering of wage contracts.

4 A Model of Uneven Staggering

This section presents a stylized macroeconomic model featuring a simple form of contractual rigidities that reproduces some of the main differences in impulse responses across quarters observed in the data. The model has three types of agents: *i*) Firms, which use labor to produce output competitively; *ii*) households, who supply labor and consume firms' output; and *iii*) a policymaker, who targets the nominal interest rate according to a Taylor-type rule.

The model's propagation mechanism in response to a temporary shock to the nominal interest rate is very simple. Firms are perfectly competitive and hence the nominal wage rigidity in the labor market is transmitted to prices. Because of this inherited rigidity, inflation does not respond immediately to changes in the nominal interest rate. Consequently, in the short run changes in the nominal interest rate translate into changes in the real interest rate and hence into changes in aggregate demand conditions. Monetary policy, however, is neutral in the long run.

4.1 Firms

Competitive firms use a continuum of differentiated labor inputs, $h(i)$, $i \in [0,1]$ to produce final output, Y, with a Dixit-Stiglitz technology:[18]

$$Y_t = \left[\int_0^1 h_t(i)^{\frac{\theta-1}{\theta}} di \right]^{\frac{\theta}{\theta-1}},$$

[18] We omit subindices for each firm and simply work with the aggregate of firms, since in equilibrium all firms are identical.

where $\theta > 1$ measures the elasticity of substitution between different types of labor, and t indicates time expressed in quarters.[19] Firms maximize profits taking prices P_t, and wages $w_t(i)$, $\forall i$, as given:

$$\max_{\{h_t(i)\}, i\in[0,1]} P_t Y_t - \int_0^1 w_t(i) h_t(i) di.$$

Firms' demand for the services of labor input i is then:

$$h_t(i) = \left(\frac{w_t(i)}{W_t}\right)^{-\theta} Y_t, \qquad i \,\epsilon[0,1], \tag{6}$$

where W_t is the aggregate wage index, defined as:

$$W_t = \left[\int_0^1 w_t(i)^{1-\theta} di\right]^{\frac{1}{1-\theta}}.$$

Because firms are perfectly competitive, in equilibrium the aggregate price level P_t will equal the aggregate wage index W_t.

4.2 Households

Infinitely-lived households derive utility from consumption and disutility from labor effort. The representative household's expected utility is given by:

$$E_t \left\{ \sum_{j=0}^{\infty} \beta^{t+j} \left[u(C_{t+j} - \eta C_{t+j-1}) - v\left(\int_0^1 h_t(i) di \right) \right] \right\}, \tag{7}$$

[19]The Dixit-Stiglitz technology, together with the assumption of differentiated labor input, provides a motivation for the assumption of workers' monopoly power. The measure of workers' monopoly power is equal to $\frac{\theta}{\theta-1}$.

where $\beta \in (0,1)$ is the intertemporal discount factor, C_t is consumption at time t, u is the period utility derived from consumption, with $u' > 0$, $u'' < 0$, $\eta \in [0,1)$ is the extent of internal habit persistence, and v is the disutility of labor effort, with $v' > 0$ and $v'' \leq 0$. Expression (7) indicates that each household offers all types of labor $h(i)$, $i \in [0,1]$. This is a convenient simplification that allows us to work with a representative agent and neglect distributional considerations.[20] For simplicity, we assume that capital markets allow for riskless borrowing and lending. Households face a budget constraint given by:

$$A_t + P_t C_t = R_{t-1} A_{t-1} + \int_0^1 w_t(i) h_t(i) di, \qquad \forall t,$$

where R_t is the riskless nominal interest rate between t and $t+1$, and A_t denotes total nominal assets held by households at time t.[21] Optimization of the sequence of consumption implies:

$$E_t(\Lambda_t P_t) = E_t \left[u_c(C_t - \eta C_{t-1}) - \beta \eta u_c(C_{t+1} - \eta C_t) \right], \tag{8}$$

where Λ_t is the marginal utility of nominal income at t, which satisfies:

$$\Lambda_t = \beta(1 + R_t) \ E_t \Lambda_{t+1}.$$

4.3 Wage Setting with Nominal Rigidities and Staggering

Households enjoy monopoly power in the labor markets and set wages in order to optimize their expected utility. If workers were allowed to change wages every period, then at each point in time

[20] For this reason, subindices for the household are omitted. An alternative specification has each household provide only one type of labor. The two models yield identical predictions if we further assume that in the latter case households can completely insure against idiosyncratic wage fluctuations. See Woodford (2003) for details.

[21] A standard no-Ponzi-game condition applies to the intertemporal version of the period budget constraint.

the wage would be given by:

$$w_t^*(i) = \arg\max_{w_t(i)} \left[\Lambda_t \int_0^1 w_t(i) h_t(i) di - v \left(\int_0^1 h_t(i) di \right) \right],$$

where $h_t(i)$ is firms' demand for type-i labor as given by equation (6). That is, the optimal wage $w_t^*(i)$ would satisfy:

$$\Lambda_t w_t(i) - v' \left(\int_0^1 h_t(i) di \right) \frac{\theta}{\theta - 1} = 0, \tag{9}$$

which simply says that the marginal utility of labor income is equal to a markup $\frac{\theta}{\theta-1}$ times the marginal disutility of work.

In our setting, however, wages are not flexible. They are instead set according to a variant of the mechanism proposed by Calvo (1983). The difference from the standard Calvo setup is that we allow for time-varying probabilities of wage changes as a way to feature the clustering of contracts at certain times of the calendar year. For simplicity and to keep a close parallel with the empirical exercise, we split the year into four quarters and assume that for any labor type the probability of resetting wages in quarter j is $(1 - \alpha_j)$, with $j = 1, ..., 4$. When a labor type receives the signal to change the 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. The optimal wage for labor-type i resetting the contract in the first quarter is then:

$$w_t^*(i) = \arg\max_{\{w_t(i)\}} E_t \left\{ \begin{array}{c} \sum_{j=0}^{\infty} \prod_{k=1}^{4} \alpha_k^j \beta^{4j} \{ \left[\Lambda_{t+4j} \int_0^1 w_t(i) h_{t+4j}(i) di - v\left(\int_0^1 h_{t+4j}(i) di \right) \right] + \\ +\alpha_2 \beta \left[\Lambda_{t+4j+1} \int_0^1 w_t(i) h_{t+4j+1}(i) di - v\left(\int_0^1 h_{t+4j+1}(i) di \right) \right] + \\ +\alpha_2 \alpha_3 \beta^2 \left[\Lambda_{t+4j+2} \int_0^1 w_t(i) h_{t+4j+2}(i) di - v\left(\int_0^1 h_{t+4j+2}(i) di \right) \right] + \\ +\alpha_2 \alpha_3 \alpha_4 \beta^3 \left[\Lambda_{t+4j+3} \int_0^1 w_t(i) h_{t+4j+3}(i) di - v\left(\int_0^1 h_{t+4j+3}(i) di \right) \right] \} \end{array} \right\} \quad (10)$$

Note, however, that this expression is valid only for wages set in the first quarter.[22] More generally, we show in the Appendix that wage maximization in any given quarter leads to the following log-linearized expression for the optimal wage:[23]

$$\tilde{w}_t^* = \frac{1 - \prod_{j=1}^{4} \alpha_j \beta^4}{\Gamma_{(q,\alpha,\beta)}} \left[\gamma \left[\left(1 + \beta\eta^2\right) \tilde{c}_t - \eta \tilde{c}_{t-1} - \beta\eta\tilde{c}_{t+1} \right] + \tilde{p}_t - \psi \tilde{h}_t \right] + \beta \sum_{k=1}^{4} q_{kt} \alpha_{k+1} \frac{\Phi_{(q,\alpha,\beta)}}{\Gamma_{(q,\alpha,\beta)}} E_t \tilde{w}_{t+1}^*, \quad (11)$$

where

$$\begin{aligned} \Gamma_{(q,\alpha,\beta)} &= 1 + \beta \{ q_{1t}\alpha_2 \left[1 + \beta\alpha_3 + \beta^2\alpha_3\alpha_4\right] + q_{2t}\alpha_3 \left[1 + \beta\alpha_4 + \beta^2\alpha_4\alpha_1\right] + \\ &\quad + q_{3t}\alpha_4 \left[1 + \beta\alpha_1 + \beta^2\alpha_1\alpha_2\right] + q_{4t}\alpha_1 \left[1 + \beta\alpha_2 + \beta^2\alpha_2\alpha_3\right] \}, \\ \Phi_{(q,\alpha,\beta)} &= 1 + \{ \beta q_{1t}\alpha_3 \left[1 + \beta\alpha_4 + \beta^2\alpha_4\alpha_1\right] + q_{2t}\alpha_4 \left[1 + \beta\alpha_1 + \beta^2\alpha_1\alpha_2\right] + \\ &\quad q_{3t}\alpha_1 \left[1 + \beta\alpha_4 + \beta^2\alpha_4\alpha_3\right] + q_{4t}\alpha_2 \left[1 + \beta\alpha_3 + \beta^2\alpha_3\alpha_4\right] \}, \end{aligned}$$

$$\gamma = \frac{1}{1-\eta} \frac{\rho}{1-\beta\eta}, \;\; \rho = -\frac{x u''(x)}{u'(x)}, \;\; \psi = -\frac{x v''(x)}{v'(x)},$$

[22] The corresponding expression for workers setting wages in the second quarter can be obtained by subtituting α_{j+1} for α_j , for $j = 1, 2, 3$ and substituting α_4 for α_1 in expression (10). Appropriate substitutions lead to the formulae for the optimal wage in the third and fourth quarters.

[23] Since all workers resetting their wages in a given quarter will choose the same wage, we can omit the i's from the expression.

$q_{jt} = q_{jt-4}$ with $j = 1, 2, 3, 4$ denote dummy variables equal to 1 when t corresponds to the jth quarter and 0 otherwise, and $\tilde{w}_t^*$, $\tilde{p}_t (= \tilde{w}_t)$, $\tilde{c}_t$, and $\tilde{h}_t$ are the log-linearized deviations around the steady state of W_t^*, $P_t (= W_t)$, C_t, and $H_t = \int_0^1 h_t(i)di$, respectively. Equation (11) says that – in log deviations from the steady state – the optimal wage is the weighted average between the optimal wage that would be chosen if wages were flexible, and the expected optimal wage in the following period, $E_t \tilde{w}_{t+1}^*$.[24]

The aggregate wage level is a weighted average between the optimal wage of the labor types that received the signal to change and the wages of those that did not get the signal. By the law of large numbers, the proportion of labor types renegotiating wages at t will be equal to $(1 - \alpha_j)$. Therefore, the quarter-dependent law of motion for the aggregate wage in log deviations from the steady state is given by:

$$\tilde{w}_t = \left[1 - \sum_{k=1}^{4} q_{kt}\alpha_k\right] \tilde{w}_t^* + \sum_{k=1}^{4} q_{kt}\alpha_k \tilde{w}_{t-1}. \quad (12)$$

4.4 Monetary Policy

We assume that the monetary authority follows a Taylor-type reaction function:

$$\tilde{\imath}_t = \delta_p \tilde{\pi}_t + \delta_y \tilde{y}_t + \xi \tilde{\imath}_{t-1} + v_t, \quad (13)$$

where $\tilde{\imath}_t = \ln R_t - \ln \bar{R}$ is the nominal interest rate (in deviation from its steady state level), $\tilde{\pi}_t = \tilde{p}_t - \tilde{p}_{t-1}$ is the inflation rate, and $\tilde{y}_t = \ln Y_t - \ln \bar{Y}$ is the output gap. In order to ensure the existence of a solution, we impose the restrictions $\frac{\delta_p}{1-\xi} > 1$, $\delta_y > 0$ and $\xi \in (0, 1)$ on the parameters in the reaction

[24]Note that the weights add to 1:

$$\frac{1 - \prod \alpha_j \beta^4}{\Gamma(q, \alpha, \beta)} + \beta \left(\sum q_{kt}\alpha_{k+1}\right) \frac{\Phi(q, \alpha, \beta)}{\Gamma(q, \alpha, \beta)} = 1.$$

If the probabilities of not changing wages are identical across quarters, $(\alpha_j = \alpha \in (0, 1)\ \forall j)$, the standard Calvo setup applies.

function. The term v_t is the policy shock whose effect we want to evaluate.

4.5 Model Solution and Calibration

The model is closed by the resource constraint, $\tilde{h}_t = \tilde{y}_t = \tilde{c}_t$. This condition, together with the log-linear Euler equation:

$$\gamma E_t \left[\left(1+\beta\eta^2\right)(\tilde{c}_{t+1}-\tilde{c}_t) - \eta(\tilde{c}_t - \tilde{c}_{t-1}) - \beta\eta(\tilde{c}_{t+2}-\tilde{c}_{t+1}) \right] = E_t(\tilde{\imath}_t - \tilde{\pi}_{t+1}), \tag{14}$$

the wage-dynamics equations (11) and (12), the Taylor-reaction function (13), and the zero-profit condition $\tilde{w}_t = \tilde{p}_t$, characterizes the model dynamics. Because of the presence of the time-varying indicators q_{jt} in the equations describing the wage dynamics, the system is non-linear. To solve the model, we use the non-linear algorithm proposed by Fuhrer and Bleakley (1996).

Table 2 summarizes the benchmark values used to calibrate the model. The discount factor, β, is set at $\beta = 0.95$.[25] The Taylor-rule parameters δ_p, δ_y, and ξ are standard in the literature, and are taken from Fuhrer (1994). The parameter ψ, which measures the degree of curvature of the function $v(.)$, the disutility of labor, that is, $\psi = -h \cdot v''/v'$, is set at $\psi = -0.10$. A value of ψ equal to 0, which is a usual benchmark, corresponds to the linear case in which the income and substitution effects of wages on labor supply cancel out; $\psi < 0$ implies that the substitution effect dominates. Finally, we calibrate the parameter governing the degree of habit persistence η at 0.3. This parameter value, along with the other calibrated parameters, yields to a reduced-form for consumption that exhibits approximately the same degree of autocorrelation we observe in macro data.

In calibrating the values for the probabilities of not changing wage in a given period, we consider

[25] Standard calibrations set this parameter at 0.99. However, recent estimates by Gali and Gertler (1999) suggest much smaller values. While our choice is a compromise, results are not sensitive to changing the parameter to either 0.99 or 0.85.

two cases. The first case corresponds to the Calvo setup with uniform staggering. In such a setup the timing of the monetary shock is irrelevant, since the probabilities of wage change across quarters are identical. We set $\alpha_j = \alpha = 0.6$ as a reference, implying that the average frequency of wage/price changes is 2.6 (=1/(1-0.6)) quarters, or eight months. The second case assumes that $\alpha_4 < \alpha_3 < \alpha_1 < \alpha_2$, i.e., it assumes uneven staggering. This corresponds to a situation in which most wages are changed during the course of the fourth and third quarters, some are changed in the first quarter, and fewer during the second quarter. This assumption is in line with anecdotal evidence suggesting that wages are reset in the later months or at the very beginning of the calendar year, with fewer changes taking place during the second quarter. In this scenario, we consider the effect of a policy shock of equal magnitude in each quarter. To keep the results comparable, we set the α_j's so that $\prod \alpha_j = \alpha^4 = 0.6^4$, that is, the probability for a given labor type of not changing wage in a given year is the same as in the Calvo setup. As discussed in the introduction, there is no systematic empirical evidence pointing to particular values for the α_j's. We calibrate the parameters as follows: $\alpha_1 = 0.7$; $\alpha_2 = 0.9$; $\alpha_3 = 0.5$; $\alpha_4 = 0.4$. The conditional frequencies implied by these parameter values for α_j's are consistent with evidence from the New England firms surveyed for the Federal Reserve System's *Beige Book*. Note that the number of firms setting wages in the jth quarter is $(1 - \alpha_j)N$, where N is the total number of firms. Thus, the proportion, n_j, of firms changing wages in the jth quarter relative to the total number of changes in a given year ($\sum(1 - \alpha_j)N$) is

$$n_j = (1 - \alpha_j)/(4 - \sum_{j=1}^{4} \alpha_j) \tag{15}$$

Our parametrization, hence, implies that 20 percent ($n_1 = (1 - \alpha_1)/(4 - \sum \alpha_j) = 0.20$) of the wage changes take place in the first quarter, 7 percent ($n_2 = (1 - \alpha_2)/(4 - \sum \alpha_j) = 0.07$) in the second quarter, 33 percent ($n_3 = (1 - \alpha_3)/(4 - \sum \alpha_3) = 0.33$) in the third quarter and 40 percent

($n_4 = (1 - \alpha_4)/(4 - \sum \alpha_j) = 0.40$) in the fourth quarter.

The results of the simulations are displayed in Figures 18 through 22, which show the impulse responses of output, wages (prices), the optimal wage, and the nominal interest rate to a temporary shock in the nominal interest rate. To make the results comparable to the identifying assumption underlying our empirical exercise, the shock occurs at the very end of period $t-1$, when all other $t-1$ variables have been already set. In the case of uniform staggering (Figure 18), the timing of the shock is irrelevant and the impulse responses can be interpreted as an average over the four quarters. In the case of uneven staggering, we distinguish between shocks in the first, second, third, and fourth quarters.

Monetary shocks in the first and second quarters (Figures $19-20$) produce a large impact on output, since *a*) few people are allowed to reset their wages optimally and *b*) those who adjust optimally do so to a lesser extent because of strategic complementarities in wage-setting decisions. The effect on output tends to be more persistent when the shock takes place in the first quarter, as wages remain fixed for longer. Monetary shocks in the third and fourth quarters have, overall, a smaller impact on output (Figures $21-22$). This small effect, however, is more persistent in the fourth quarter because the adjustment is followed by two consecutive quarters of relatively high wage rigidity. As for wages, the response is faster and larger on impact than the response to shocks that occur in the first or second quarter.

Overall, this modified Calvo setup is able to produce impulse responses to monetary shocks for output in different quarters that are broadly consistent with the empirical patterns we have documented. Standard extensions of the model that allow for adjustment costs and delivery lags à la Rotemberg and Woodford (1997), for example, should refine the matching between our simple DSGE model and the impulse-response functions implied by the data. In particular, these extensions would allow for a lagged response of prices and output to monetary policy shocks. We leave this, and other potential extensions of the seasonal dependent model, for future work.

5 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 accentuated 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/or fourth quarters of the year. The conventional finding that monetary shocks affect output with long delays and that the effect is highly persistent results from the aggregation of these different output impulse responses. Moreover, we argue that the differential responses are not driven by different types of monetary shocks nor by different "states" of the economy across quarters.

Encouraged by anecdotal evidence on the timing of wage changes, we propose a potential explanation for the differential responses based on contractual lumping and develop a theoretical DSGE model whose impulse responses broadly mimic those found in the data. Expanding the model to allow for adjustment costs and information lags should improve the ability of the model to match other features of the empirical impulse responses. Of course, we cannot definitively claim that time-varying contractual rigidity is the ultimate explanation for the differences in impulse-responses across quarters. Other seasonal factors may be at play and this should make an interesting subject for future research.

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A Appendix

In this section, we discuss in more detail the conditions that lead to equation (11). As argued before, for labor-type i setting wages in the first quarter, $w_t^*(i)$ satisfies:

$$w_t^*(i) = \arg\max_{\{w_t(i)\}} E_t \left\{ \begin{array}{c} \sum_{j=0}^{\infty} \prod_{i=1}^{4} \alpha_i^j \beta^{4j} \left[\Lambda_{t+4j} \int_0^1 w_t(i) h_{t+4j}(i) di - v\left(\int_0^1 h_{t+4j}(i) di \right) \right] + \\ + \sum_{j=0}^{\infty} \alpha_2 \prod_{i=1}^{4} \alpha_i^j \beta^{4j+1} \left[\Lambda_{t+4j+1} \int_0^1 w_t(i) h_{t+4j+1}(i) di - v\left(\int_0^1 h_{t+4j+1}(i) di \right) \right] \\ + \sum_{j=0}^{\infty} \alpha_2 \alpha_3 \prod_{i=1}^{4} \alpha_i^j \beta^{4j+2} \left[\Lambda_{t+4j+2} \int_0^1 w_t(i) h_{t+4j+2}(i) di - v\left(\int_0^1 h_{t+4j+2}(i) di \right) \right] \\ + \sum_{j=0}^{\infty} \alpha_2 \alpha_3 \alpha_4 \prod_{i=1}^{4} \alpha_i^j \beta^{4j+3} \left[\Lambda_{t+4j+3} \int_0^1 w_t(i) h_{t+4j+3}(i) di - v\left(\int_0^1 h_{t+4j+3}(i) di \right) \right] \end{array} \right\}$$

That is:

$$E_t \left\{ \begin{array}{c} \sum_{j=0}^{\infty} \prod_{i=1}^{4} \alpha_i^j \beta^{4j} \left(\frac{w_t(i)}{W_{t+4j}} \right)^{-\theta} L_{t+4j} \{ \left[\Lambda_{t+4j} w_t(i) + v'\left(\int_0^1 h_{t+4j}(i) di \right) \frac{\theta}{1-\theta} \right] + \\ \alpha_2 \beta \left(\frac{w_t(i)}{W_{t+4j+1}} \right)^{-\theta} L_{t+4j+1} \left[\Lambda_{t+4j+1} w_t(i) + v'\left(\int_0^1 h_{t+4j+1}(i) di \right) \frac{\theta}{1-\theta} \right] + \\ \alpha_2 \alpha_3 \beta^2 \left(\frac{w_t(i)}{W_{t+4j+2}} \right)^{-\theta} L_{t+4j+2} \left[\Lambda_{t+4j+2} w_t(i) + v'\left(\int_0^1 h_{t+4j+2}(i) di \right) \frac{\theta}{1-\theta} \right] + \\ \alpha_2 \alpha_3 \alpha_4 \beta^3 \left(\frac{w_t(i)}{W_{t+4j+3}} \right)^{-\theta} L_{t+4j+3} \left[\Lambda_{t+4j+3} w_t(i) + v'\left(\int_0^1 h_{t+4j+3}(i) di \right) \frac{\theta}{1-\theta} \right] \} + \end{array} \right\} = 0$$

To keep the analysis simple, let us assume for the moment that period t corresponds to a first quarter (and hence, so do $t+4j$, $\forall j$) while $t+1+4j$, $\forall j$ corresponds to the second quarter, $t+2+4j$, $\forall j$ corrresponds to the third, and $t+3+4j$, $\forall j$ corresponds to the fourth quarter. Since all labor-types resetting their wages at a given quarter will choose the same wage, we can get rid of the i's and simply refer to the optimal wage in period t as w_t^*. The expression for the optimal wage in the first quarter becomes:

$$w_t^* \cdot E_t\{ \sum_{j=0}^{\infty} \prod_{i=1}^{4} \alpha_i^j \beta^{4j} \cdot [W_{t+4j}^{\theta} Y_{t+4j} \Lambda_{t+4j} + \alpha_2 \beta W_{t+4j+1}^{\theta} Y_{t+4j+1} \Lambda_{t+4j+1} +$$

$$+\alpha_2 \alpha_3 \beta^2 W_{t+4j+2}^{\theta} Y_{t+4j+2} \Lambda_{t+4j+2} + \alpha_2 \alpha_3 \alpha_4 \beta^3 W_{t+4j+3}^{\theta} Y_{t+4j+3} \Lambda_{t+4j+3}]\} =$$

$$= \tfrac{\theta}{1-\theta} E_t\{ \sum_{j=0}^{\infty} \prod_{i=1}^{4} \alpha_i^j \beta^{4j} \cdot [W_{t+4j}^{\theta} Y_{t+4j} v'(H_{t+4j}) + \alpha_2 \beta W_{t+4j+1}^{\theta} Y_{t+4j+1} v'(H_{t+4j+1}) +$$

$$+\alpha_2 \alpha_3 \beta^2 W_{t+4j+2}^{\theta} Y_{t+4j+2} v'(H_{t+4j+2}) + \alpha_2 \alpha_3 \alpha_4 \beta^3 W_{t+4j+3}^{\theta} Y_{t+4j+3} v'(H_{t+4j+3})]\},$$

where $H_t = \int_0^1 h_t(i)di$. Log-linearizing expression this expression around the steady state values and using $P_t = W_t$, together with the log-linearized Euler condition,

$$\tilde{p}_t \Lambda_t = \frac{-1}{1-\eta} \frac{\rho}{1-\beta\eta} \left[\left(1+\beta\eta^2\right) \tilde{c}_t - \eta \tilde{c}_{t-1} - \beta\eta \tilde{c}_{t+1} \right],$$

where $\rho = -\frac{xu''(x)}{u'}$, the optimal first-quarter wage (in log-deviations from steady state) is:

$$\begin{aligned}
\tilde{w}_t^* = \tfrac{1}{}E_t\{ \sum_{j=0}^{\infty} \prod_{i=1}^{4} \alpha_i^j \beta^{4j} \{ & \left[\gamma\left[\left(1+\beta\eta^2\right)\tilde{c}_{t+4j} - \eta\tilde{c}_{t+4j-1} - \beta\eta\tilde{c}_{t+4j+1}\right] + \tilde{w}_{t+4j} - \psi\tilde{h}_{t+4j}\right] + \\
& +\alpha_2\beta\left[\gamma\left[\left(1+\beta\eta^2\right)\tilde{c}_{t+4j+1} - \eta\tilde{c}_{t+4j} - \beta\eta\tilde{c}_{t+4j+2}\right] + \tilde{w}_{t+4j+1} - \psi\tilde{h}_{t+4j+1}\right] + \\
& +\alpha_2\alpha_3\beta^2\left[\gamma\left[\left(1+\beta\eta^2\right)\tilde{c}_{t+4j+2} - \eta\tilde{c}_{t+4j+1} - \beta\eta\tilde{c}_{t+4j+3}\right] + \tilde{w}_{t+4j+2} - \psi\tilde{h}_{t+4j+2}\right] + \\
& +\alpha_2\alpha_3\alpha_4\beta^3\left[\gamma\left[\left(1+\beta\eta^2\right)\tilde{c}_{t+4j+3} - \eta\tilde{c}_{t+4j+2} - \beta\eta\tilde{c}_{t+4j+4}\right] + \tilde{w}_{t+4j+3} - \psi\tilde{h}_{t+4j+31}\right] \}\},
\end{aligned} \tag{16}$$

where $= \sum_{j=0}^{\infty} \prod_{i=1}^{4} \alpha_i^j \beta^{4j} (1 + \alpha_2\beta + \alpha_2\alpha_3\beta^2 + \alpha_2\alpha_3\alpha_4\beta^3) = \frac{(1+\alpha_2\beta+\alpha_2\alpha_3\beta^2+\alpha_2\alpha_3\alpha_4\beta^3)}{1-\alpha_1\alpha_2\alpha_3\alpha_4\beta^4}$ and $\gamma = \frac{1}{1-\eta}\frac{\rho}{1-\beta\eta}$, $\rho = -\frac{xu''(x)}{u'(x)}$, and $\psi = -\frac{xv''(x)}{v'(x)}$

We can write $\tilde{w}_t^*$ in recursive form as:

$$\tilde{w}_t^* = \frac{1-\prod_{j=1}^{4}\alpha_j\beta^4}{1+\beta\alpha_2(1+\beta\alpha_3+\beta^2\alpha_3\alpha_4)}\left[\gamma\left[\left(1+\beta\eta^2\right)\tilde{c}_t-\eta\tilde{c}_{t-1}-\beta\eta\tilde{c}_{t+1}\right]+\tilde{p}_t-\psi\tilde{h}_t\right]+ \\ +\beta\alpha_2\frac{1+\beta\alpha_3\left[1+\beta\alpha_4+\beta^2\alpha_4\alpha_1\right]}{1+\beta\alpha_2(1+\beta\alpha_3+\beta^2\alpha_3\alpha_4)}E_t\tilde{w}_{t+1}^* \quad (17)$$

Note that equation (17) simply says that the optimal wage, $\tilde{w}_t^*$ (in log-deviation from its steady state level), is the weighted average, with weights given by the discount factor and the corresponding probabilities of remaining fixed, of the expected optimal wages that would prevail if wages were flexible. In particular, note that if the contract lasts only one period, ($\alpha_j = 1\forall j$), the condition is simply the log-linearization of equation (9):

$$\tilde{w}_t^* = \gamma\left[\left(1+\beta\eta^2\right)\tilde{c}_t-\eta\tilde{c}_{t-1}-\beta\eta\tilde{c}_{t+1}\right]+\tilde{p}_t-\psi\tilde{h}_t$$

The aggregate wage level will be a weighted average between the optimal wage set by workers that received the signal to reoptimize wages and the wages of those who did not get the signal; for simplicity, we assume that those who do not reoptimize set their wages at the average level prevailing in the previous period. 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 $(1-\alpha_1)$:

$$\tilde{w}_t = (1-\alpha_1)\tilde{w}_t^* + \alpha_1\tilde{w}_{t-1} \quad (18)$$

Equations (17) and (18) give us the laws of motions for the first quarter. More generally, and to account for the quarterly dependent probabilities, we can use the dummy variables, q_{jt}, with $j = 1,...4$ which take on the value 1 in the jth quarter and 0 otherwise. We can then write the equations governing the wage dynamics as:

$$\tilde{w}_t^* = \frac{1 - \prod_{j=1}^{4} \alpha_j \beta^4}{\Gamma_{(q,\alpha,\beta)}} \left[\gamma \left[\left(1 + \beta\eta^2\right) \tilde{c}_t - \eta\tilde{c}_{t-1} - \beta\eta\tilde{c}_{t+1} \right] + \tilde{p}_t - \psi\tilde{h}_t \right] + \beta \sum_{k=1}^{4} q_{kt}\alpha_{k+1} \frac{\Phi_{(q,\alpha,\beta)}}{\Gamma_{(q,\alpha,\beta)}} E_t \tilde{w}_{t+1}^*, \qquad (19)$$

and

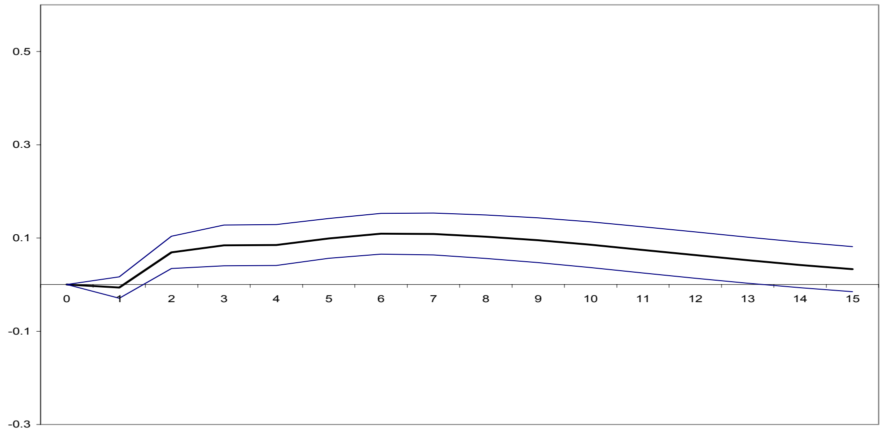
$$\tilde{w}_t = \left[1 - \left(\sum_{k=1}^{4} q_{kt}\alpha_k \right) \right] \tilde{w}_t^* + \left(\sum_{k=1}^{4} q_{kt}\alpha_k \right) \tilde{w}_{t-1}$$

where

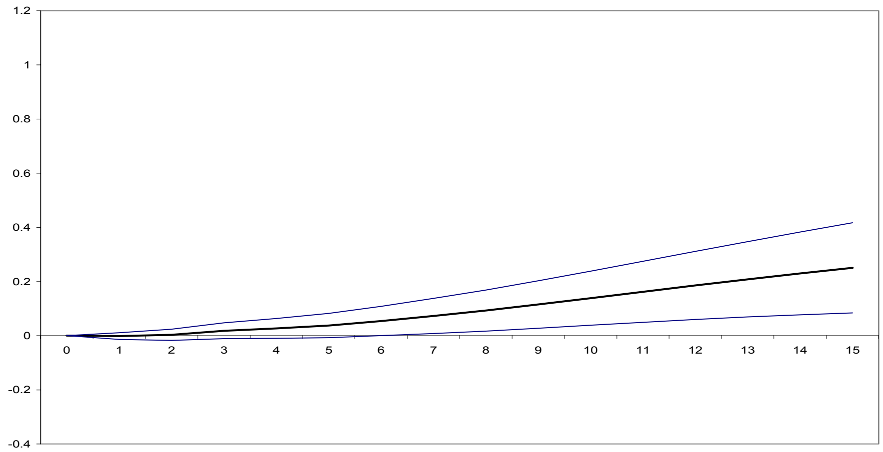
$$\begin{aligned} \Gamma_{(q,\alpha,\beta)} &= 1 + \beta\{ q_{1t}\alpha_2 \left[1 + \beta\alpha_3 + \beta^2\alpha_3\alpha_4\right] + q_{2t}\alpha_3 \left[1 + \beta\alpha_4 + \beta^2\alpha_4\alpha_1\right] + \\ &\quad + q_{3t}\alpha_4 \left[1 + \beta\alpha_1 + \beta^2\alpha_1\alpha_2\right] + q_{4t}\alpha_1 \left[1 + \beta\alpha_2 + \beta^2\alpha_2\alpha_3\right] \} \\ \Phi_{(q,\alpha,\beta)} &= 1 + \{ \beta q_{1t}\alpha_3 \left[1 + \beta\alpha_4 + \beta^2\alpha_4\alpha_1\right] + q_{2t}\alpha_4 \left[1 + \beta\alpha_1 + \beta^2\alpha_1\alpha_2\right] + \\ &\quad + q_{3t}\alpha_1 \left[1 + \beta\alpha_4 + \beta^2\alpha_4\alpha_3\right] + q_{4t}\alpha_2 \left[1 + \beta\alpha_3 + \beta^2\alpha_3\alpha_4\right]. \end{aligned}$$

FIGURE 1

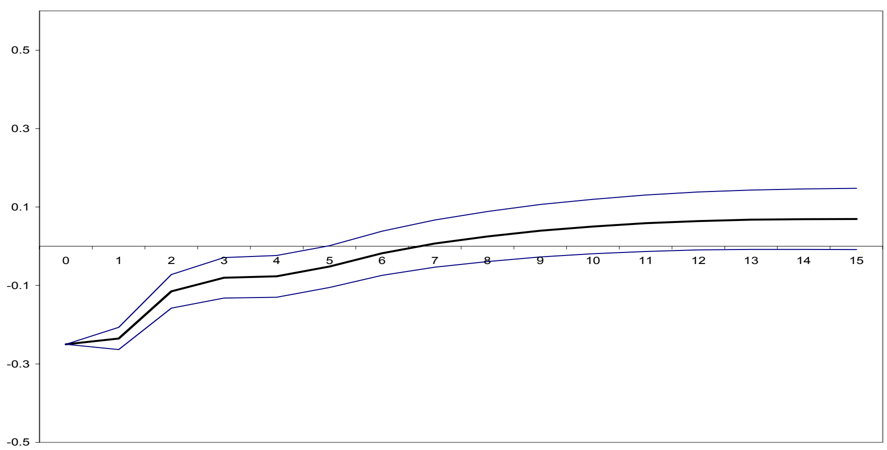
25-Basis Point Decline in Fed Funds Rate
No Quarterly Dependence 1966:Q1 to 2002:Q4



Response of GDP



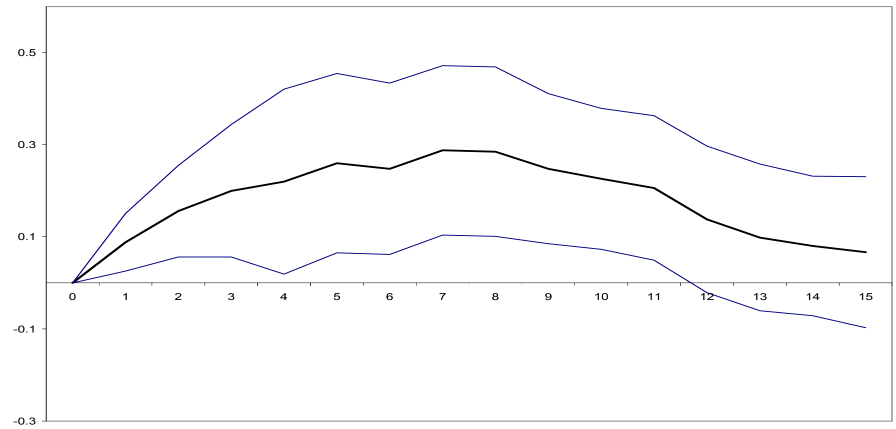
Response of PGDP



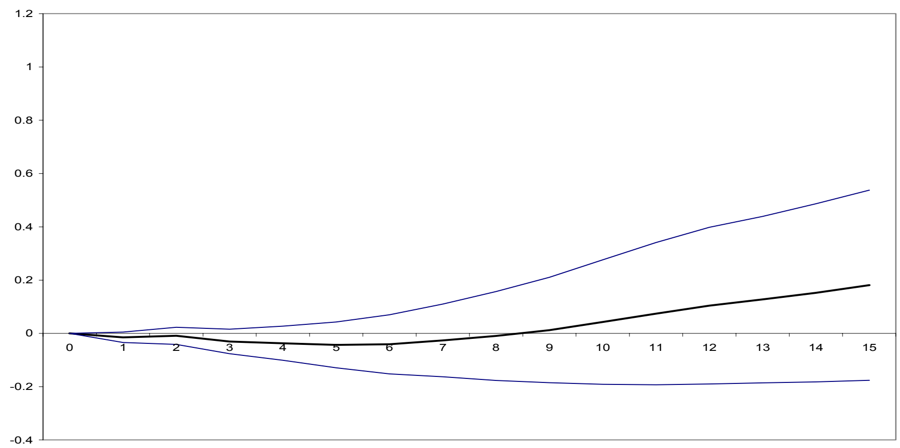
Response of FFR

FIGURE 2

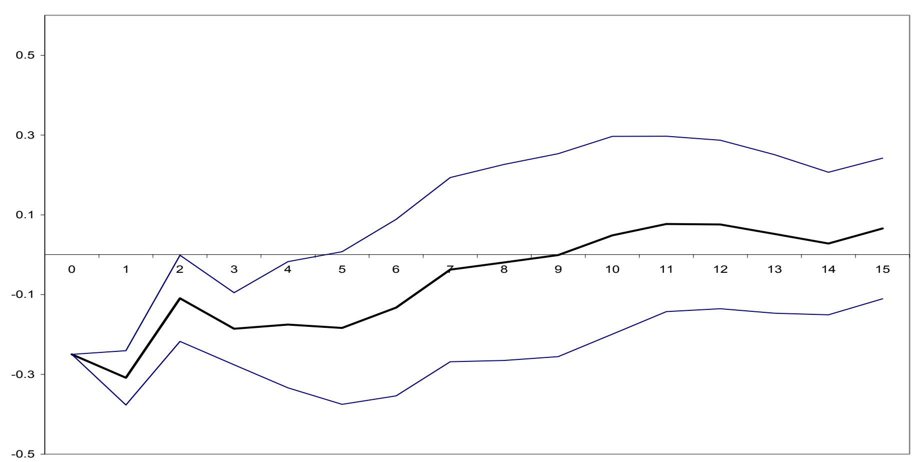
25-Basis Point Decline in Fed Funds Rate in Q1
Quarterly Dependence. Benchmark Model 1966:Q1 to 2002:Q4



Response of GDP



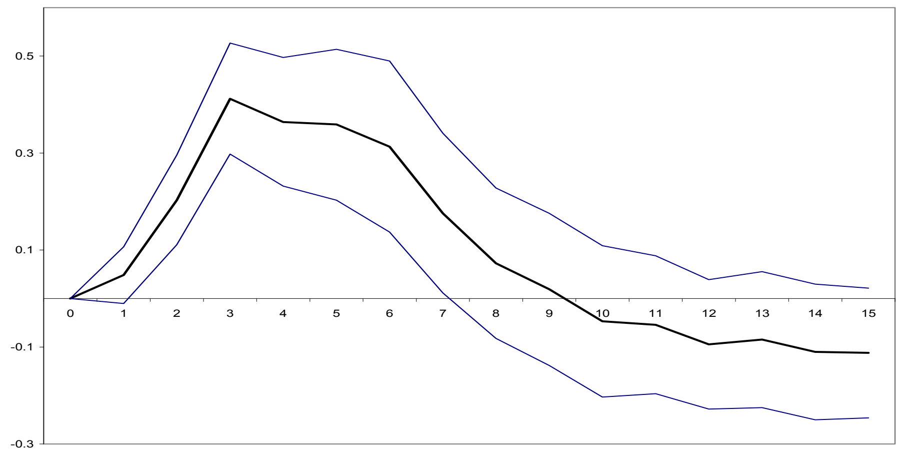
Response of PGDP



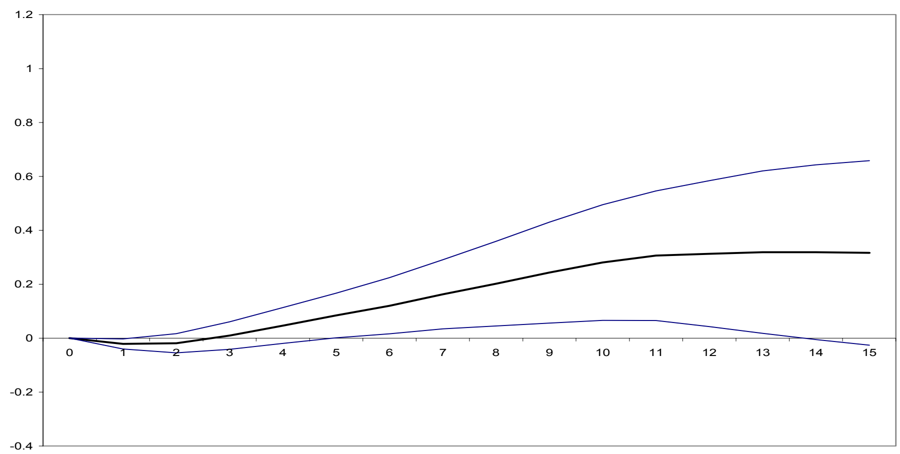
Response of FFR

FIGURE 3

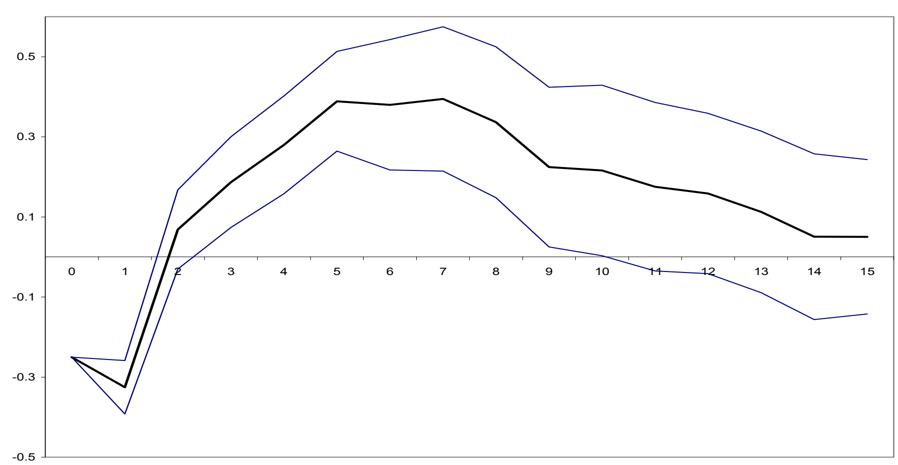
25-Basis Point Decline in Fed Funds Rate in Q2
Quarterly Dependence. Benchmark Model 1966:Q1 to 2002:Q4



Response of GDP



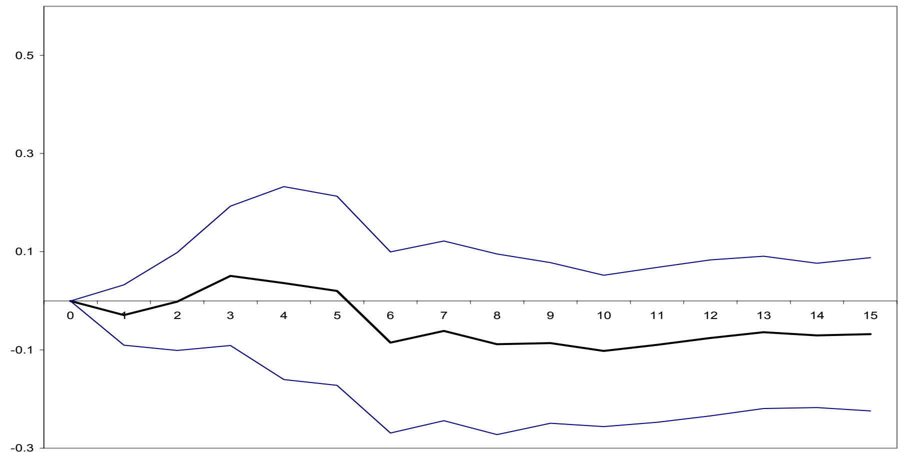
Response of PGDP



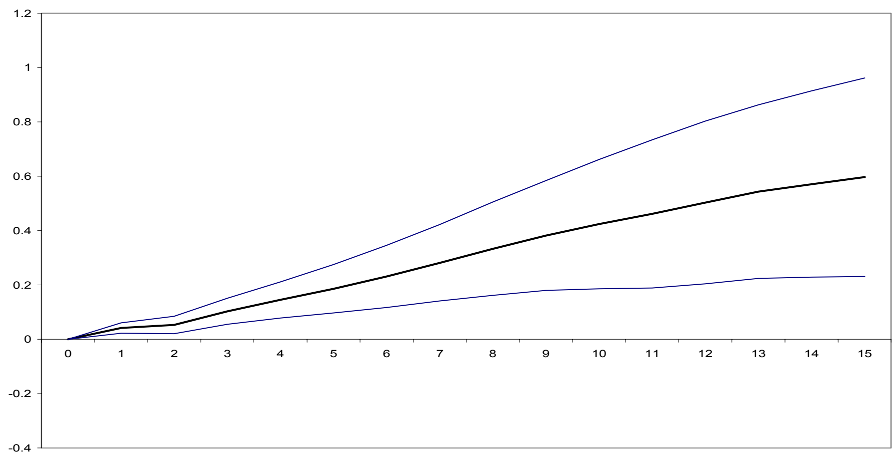
Response of FFR

FIGURE 4

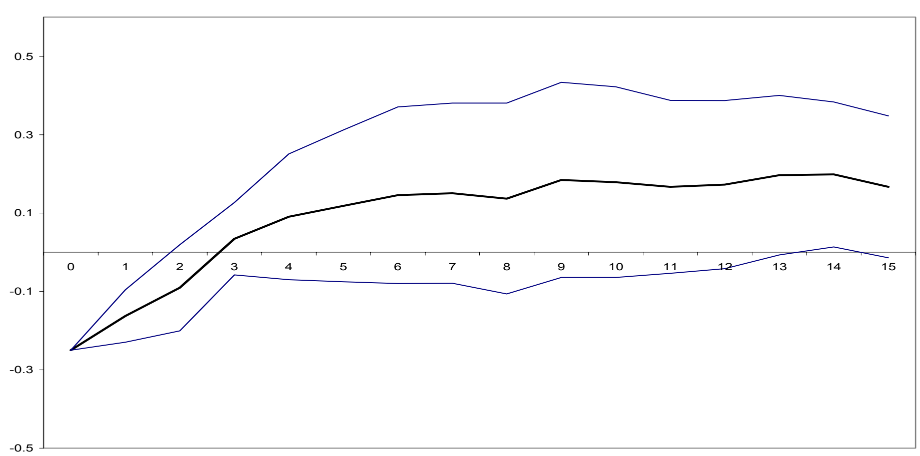
25-Basis Point Decline in Fed Funds Rate in Q3
Quarterly Dependence. Benchmark Model 1966:Q1 to 2002:Q4



Response of GDP



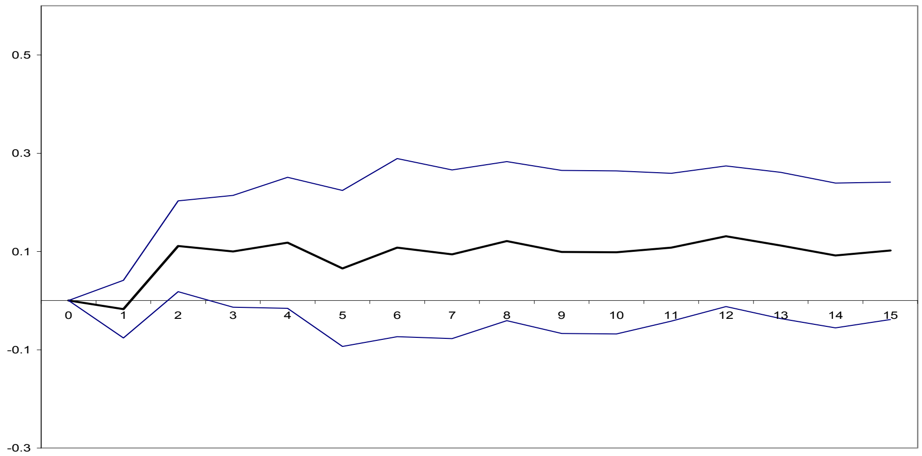
Response of PGDP



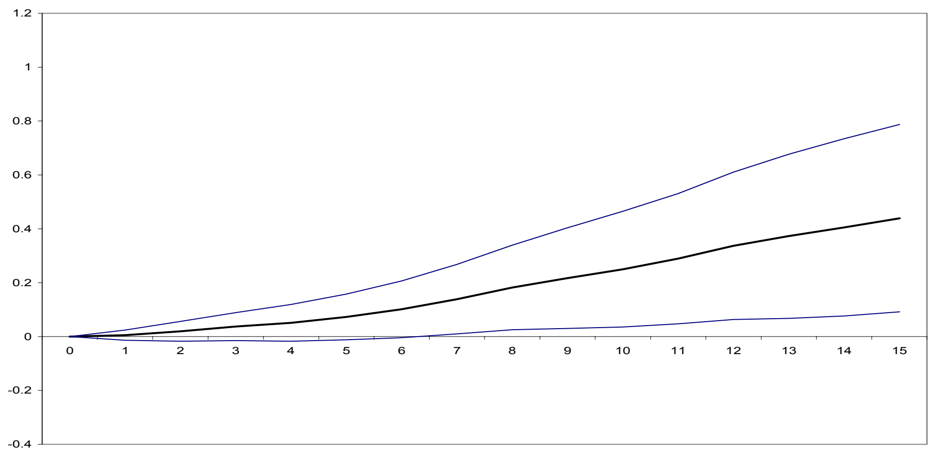
Response of FFR

FIGURE 5

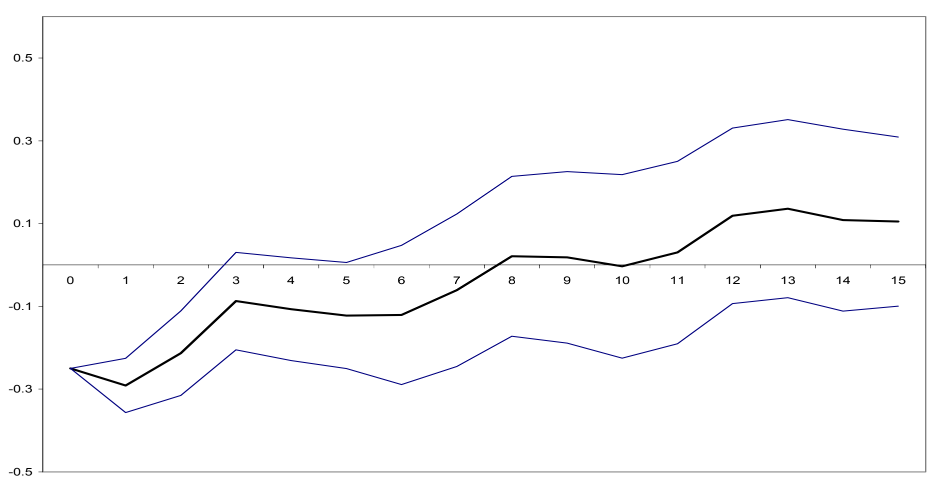
25-Basis Point Decline in Fed Funds Rate in Q4
Quarterly Dependence. Benchmark Model 1966:Q1 to 2002:Q4



Response of GDP



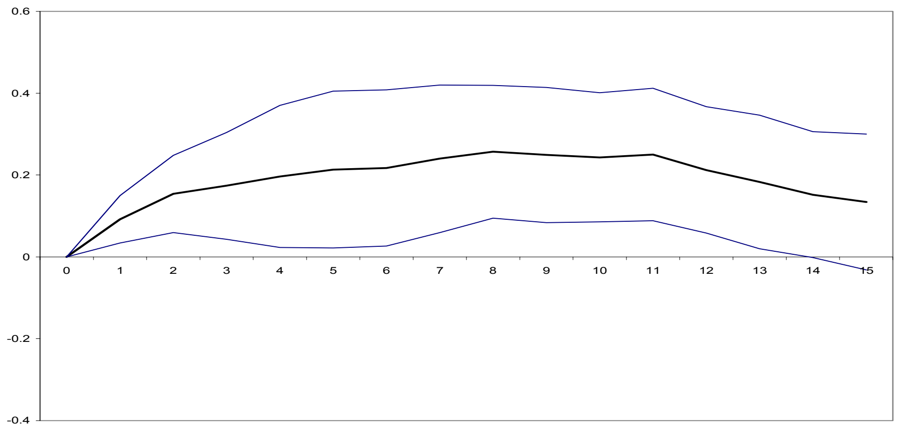
Response of PGDP



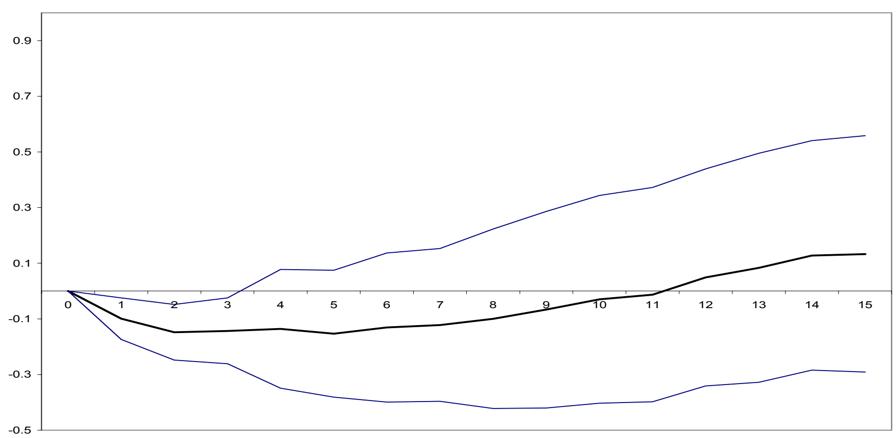
Response of FFR

FIGURE 6

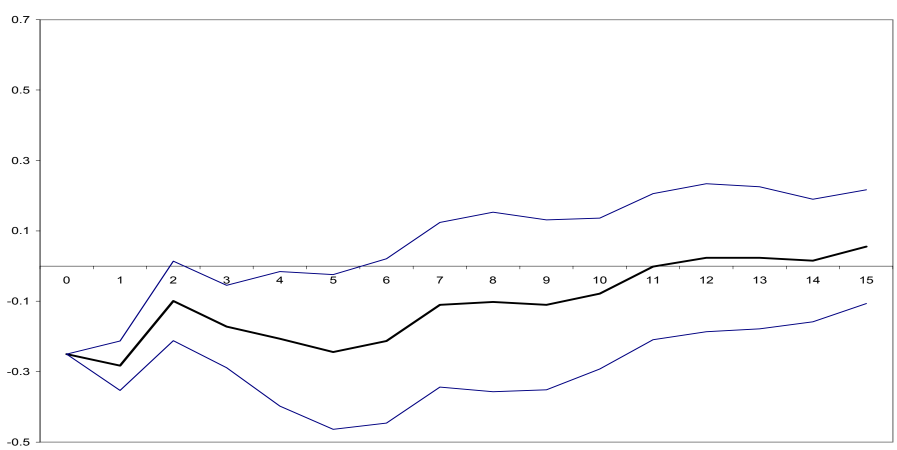
25-Basis Point Decline in Fed Funds Rate in Q1
Quarterly Dependence. Wage-System 1966:Q1 to 2002:Q4



Response of GDP



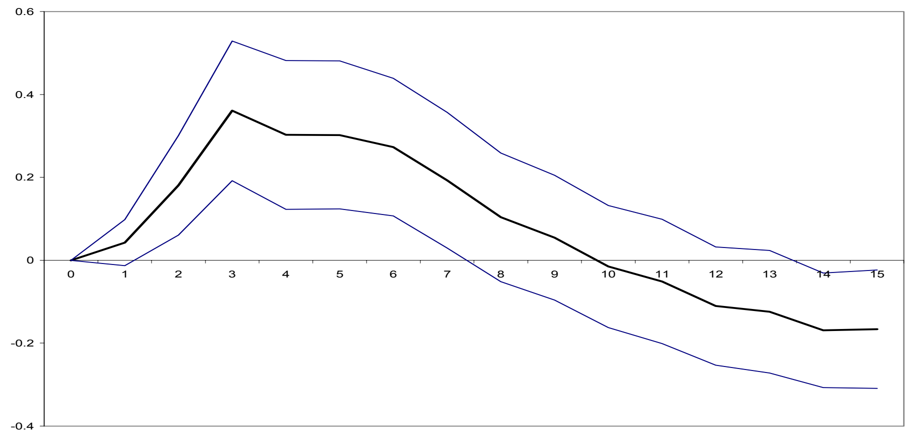
Response of Wage



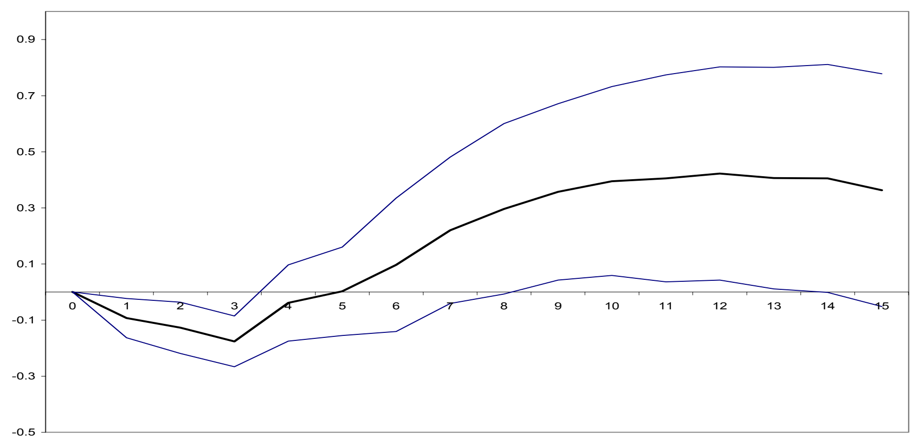
Response of FFR

FIGURE 7

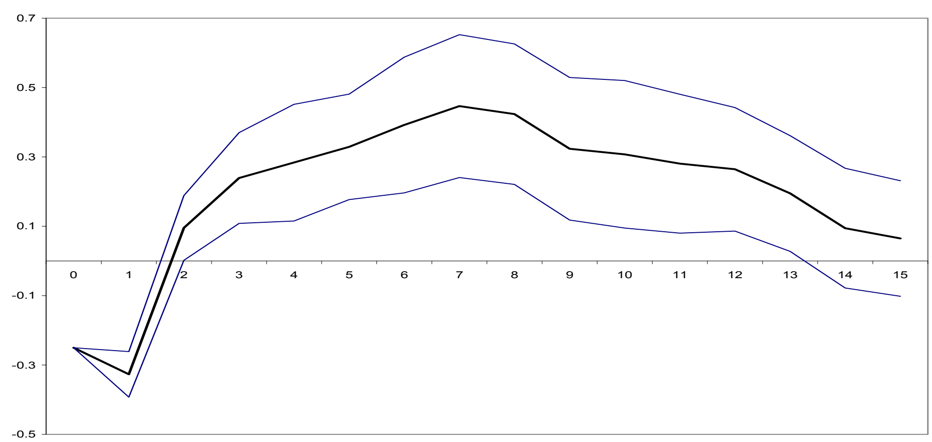
25-Basis Point Decline in Fed Funds Rate in Q2
Quarterly Dependence. Wage-System 1966:Q1 to 2002:Q4



Response of GDP



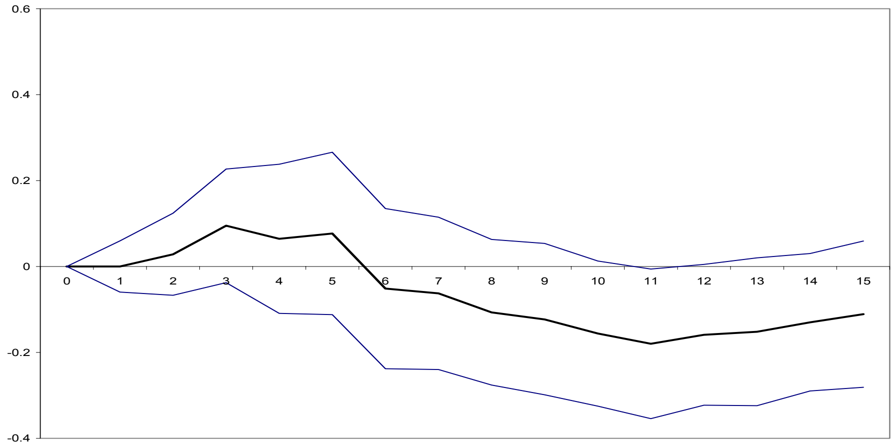
Response of Wage



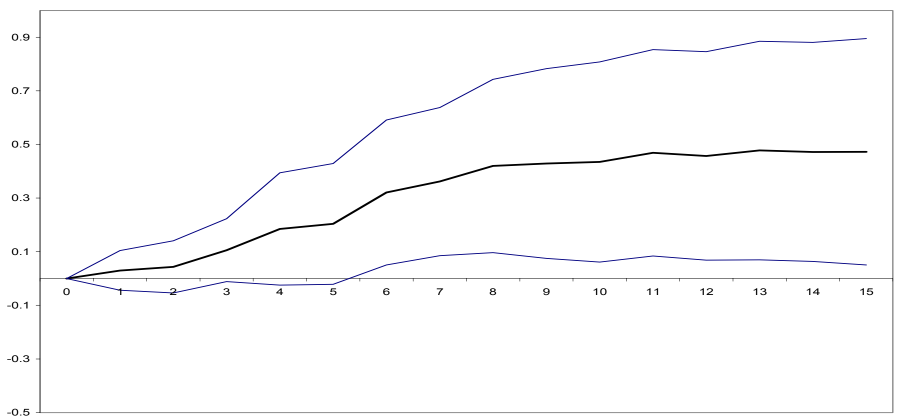
Response of FFR

FIGURE 8

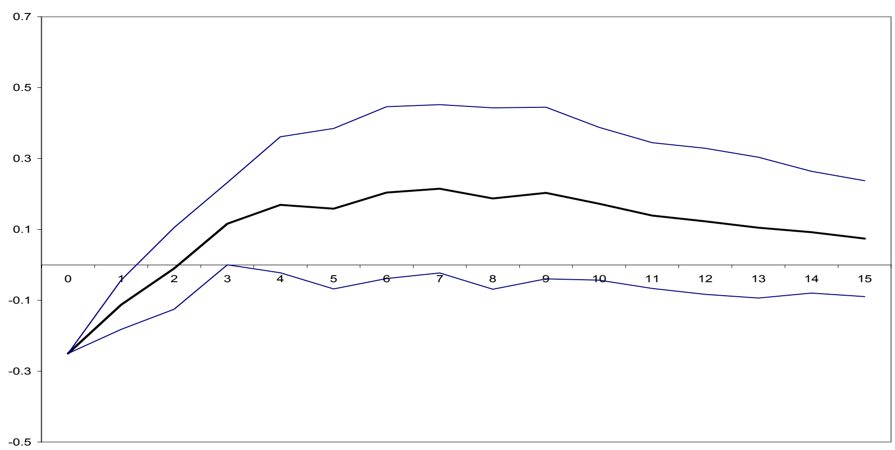
25-Basis Point Decline in Fed Funds Rate in Q3
Quarterly Dependence. Wage-System 1966:Q1 to 2002:Q4



Response of GDP



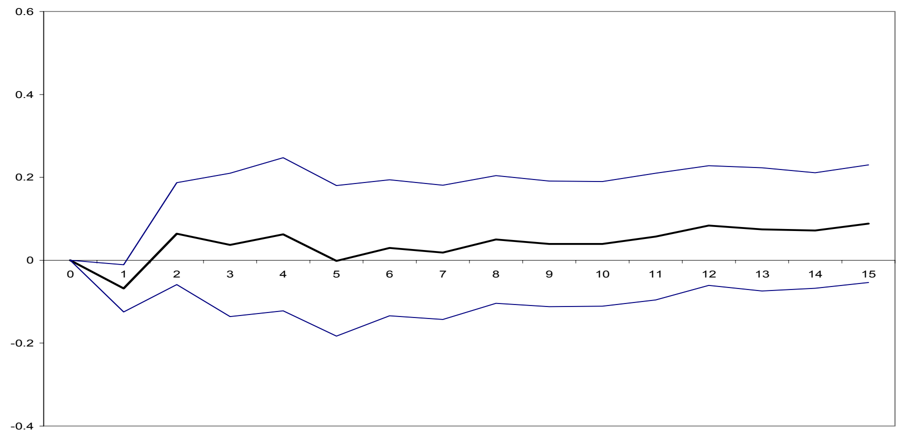
Response of Wage



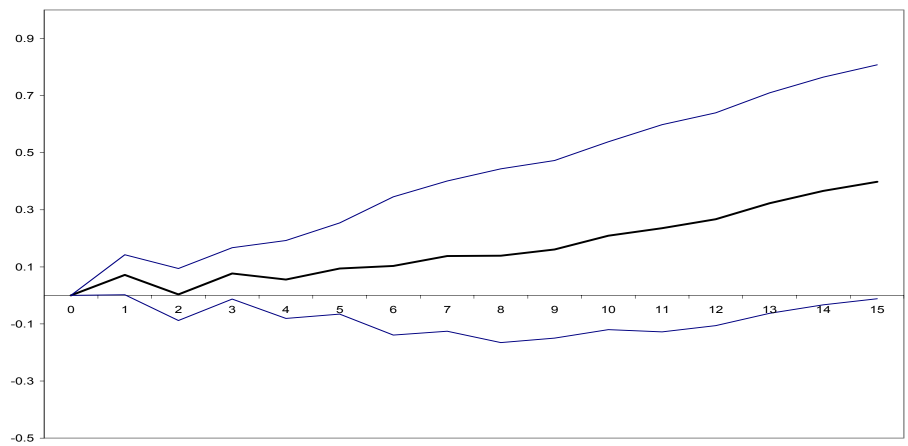
Response of FFR

FIGURE 9

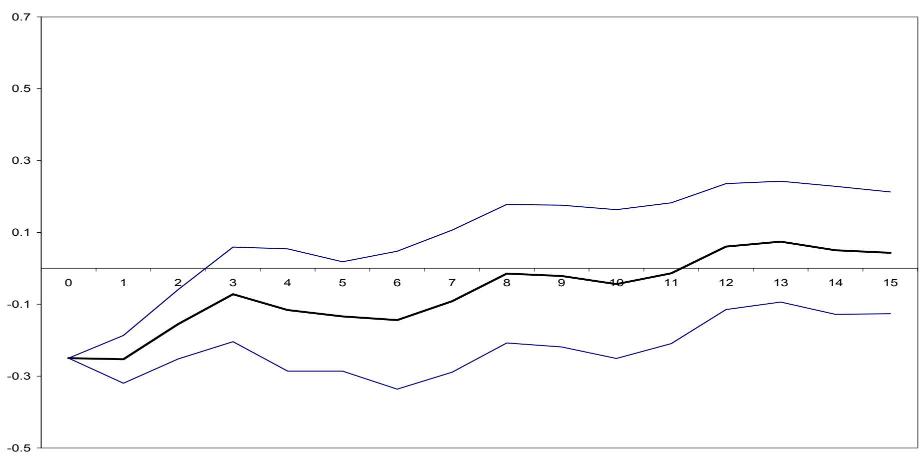
25-Basis Point Decline in Fed Funds Rate in Q4
Quarterly Dependence. Wage-System 1966:Q1 to 2002:Q4



Response of GDP



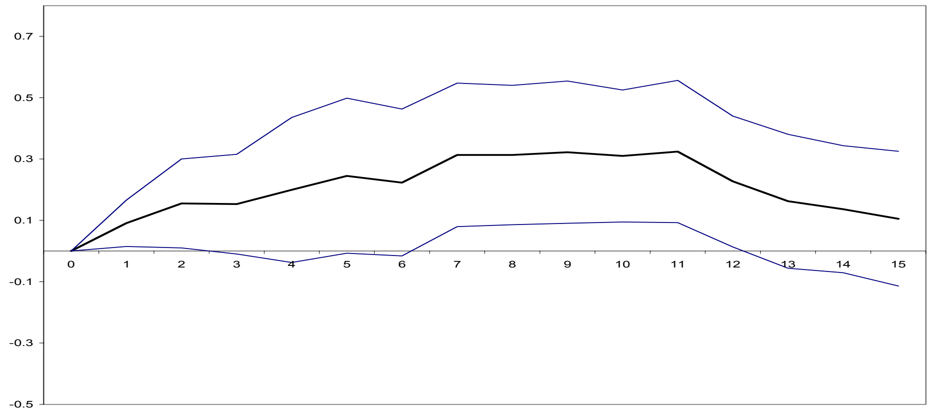
Response of Wage



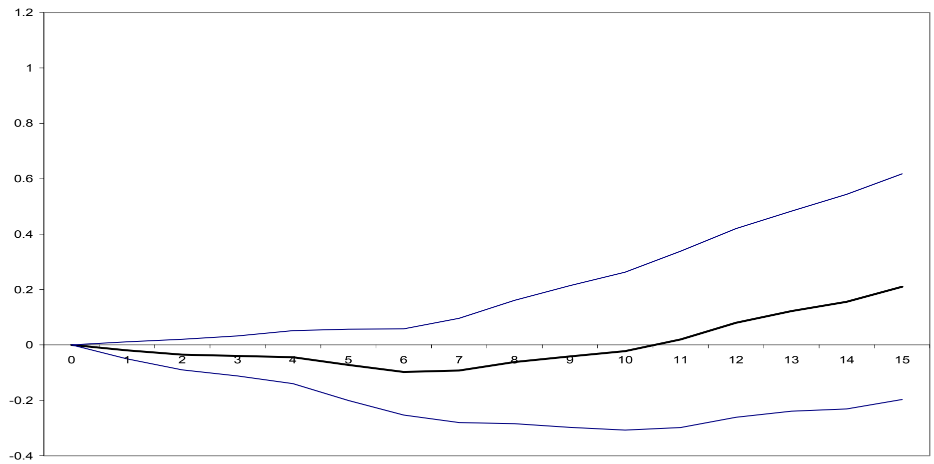
Response of FFR

FIGURE 10

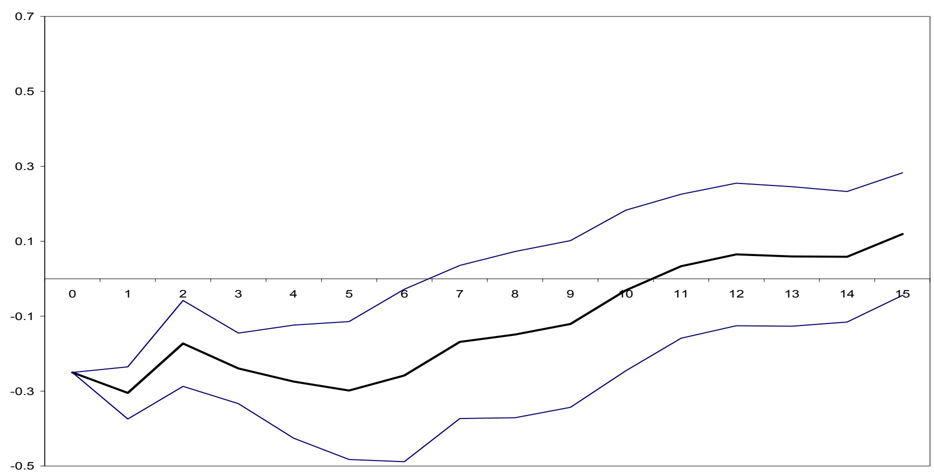
25-Basis Point Decline in Fed Funds Rate in Q1
Quarterly Dependence. NSA Data-System 1966:Q1 to 2002:Q4



Response of GDP



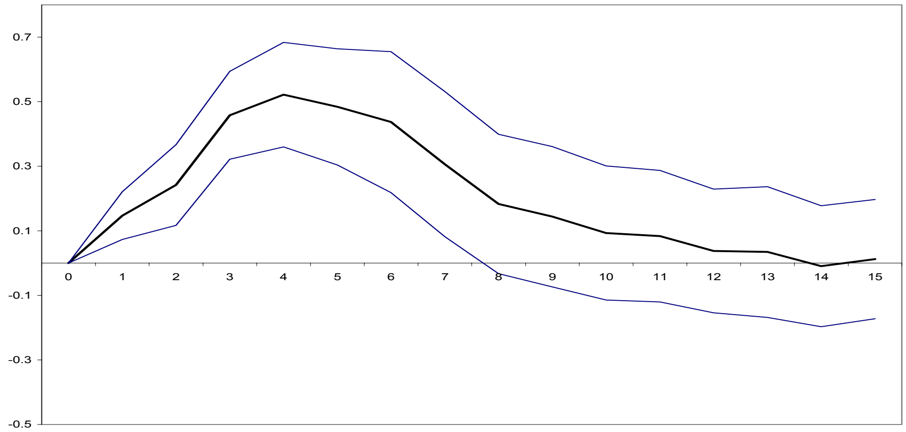
Response of CPI



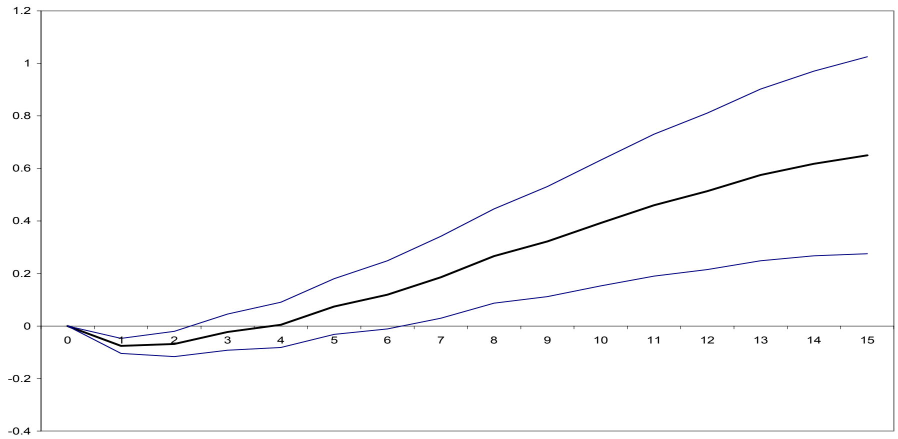
Response of FFR

FIGURE 11

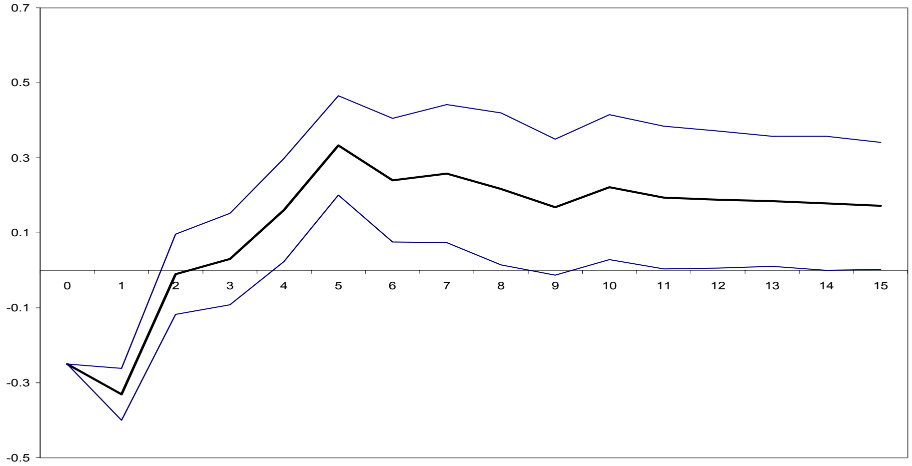
25-Basis Point Decline in Fed Funds Rate in Q2
Quarterly Dependence. NSA Data-System 1966:Q1 to 2002:Q4



Response of GDP



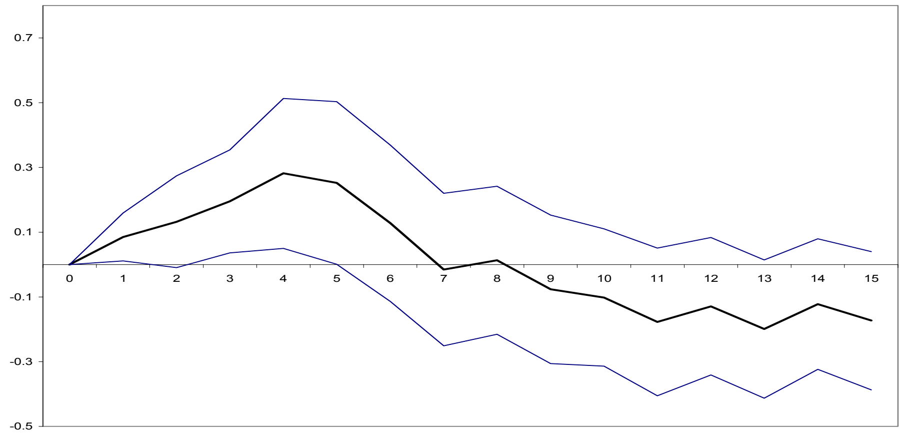
Response of CPI



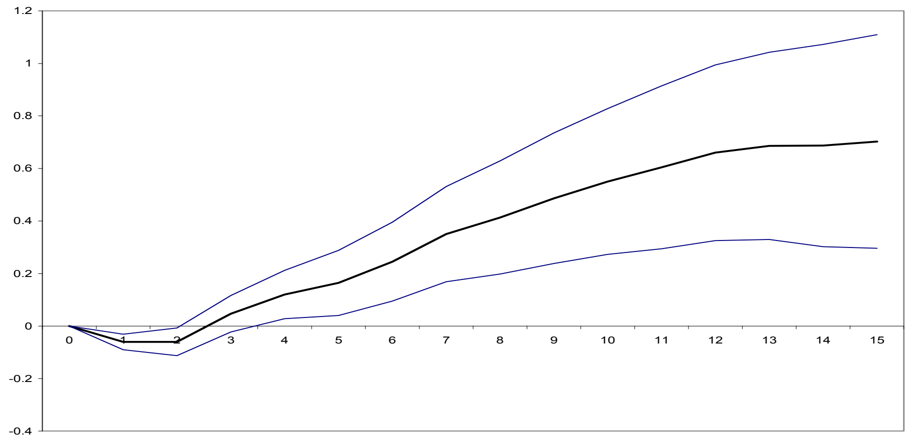
Response of FFR

FIGURE 12

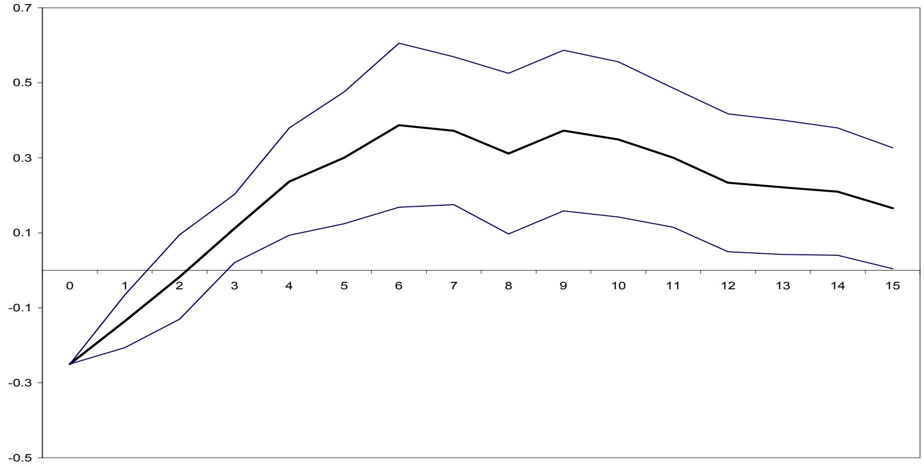
25-Basis Point Decline in Fed Funds Rate in Q3
Quarterly Dependence. NSA Data-System 1966:Q1 to 2002:Q4



Response of GDP



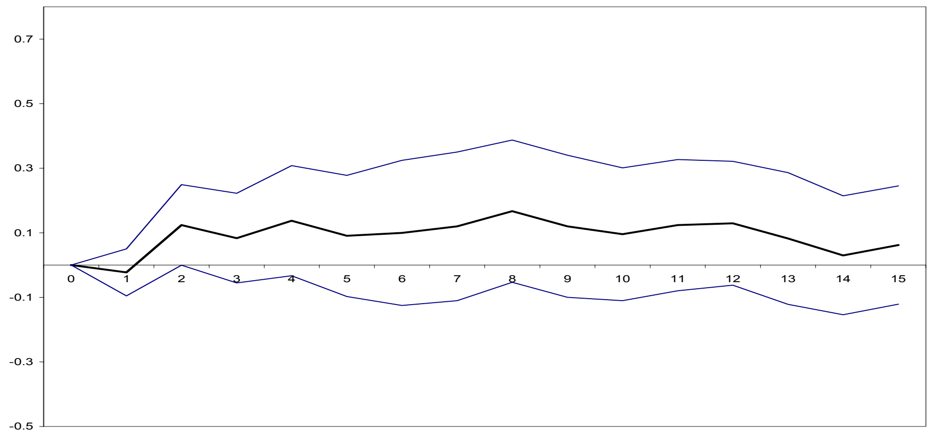
Response of CPI



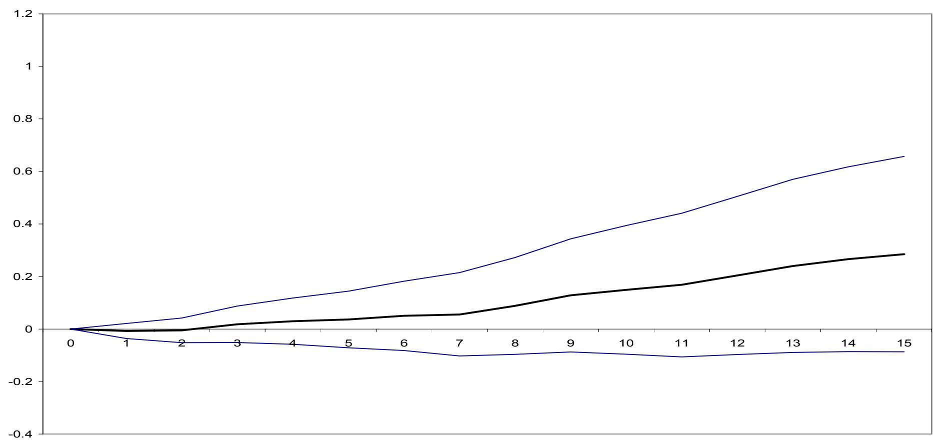
Response of FFR

FIGURE 13

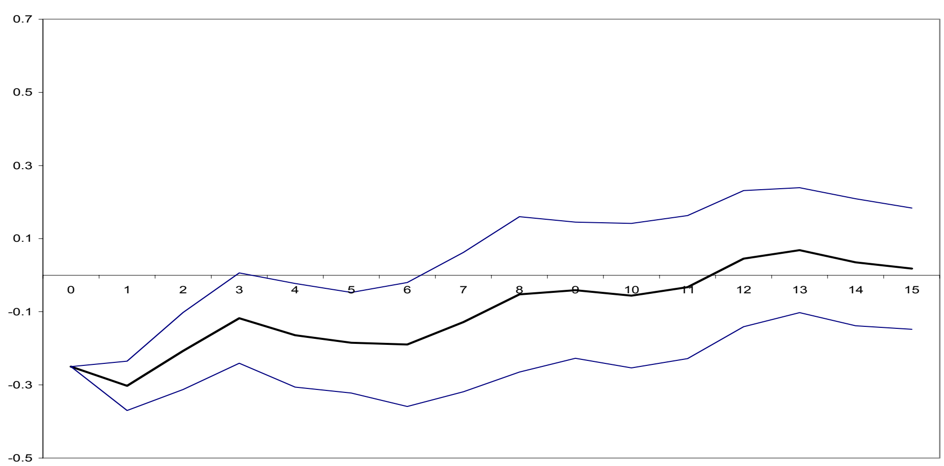
25-Basis Point Decline in Fed Funds Rate in Q4
Quarterly Dependence. NSA Data-System 1966:Q1 to 2002:Q4



Response of GDP



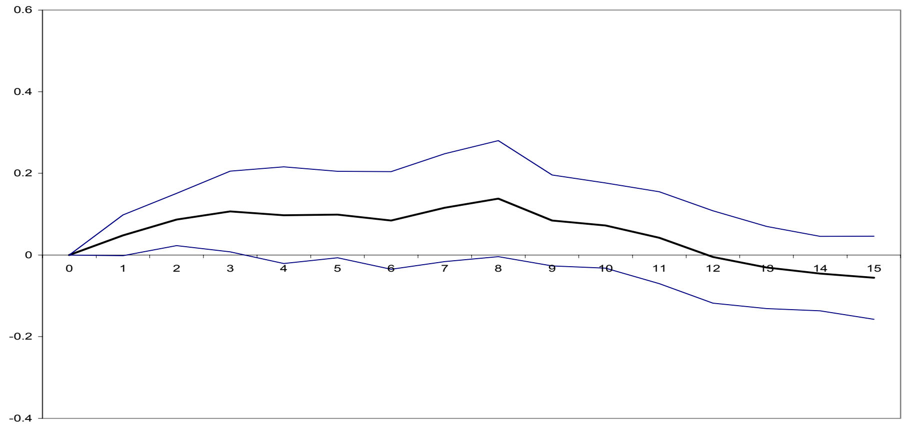
Response of CPI



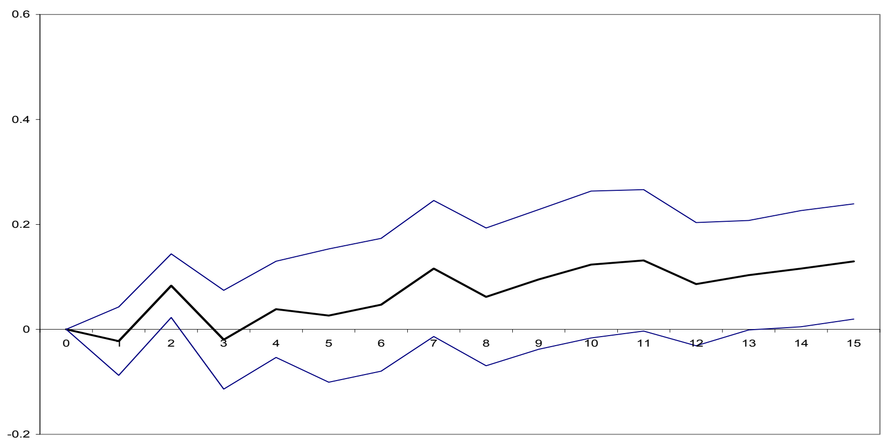
Response of FFR

FIGURE 14

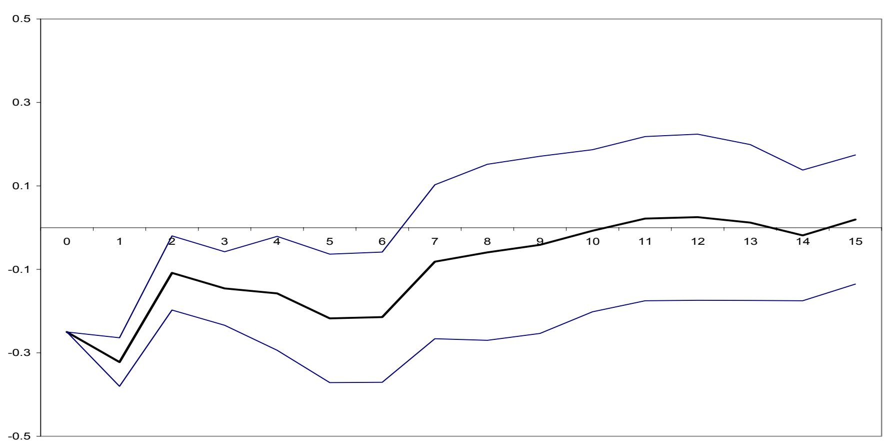
25-Basis Point Decline in Fed Funds Rate in Q1
Quarterly Dependence. Output-Gap System 1966:Q1 to 2002:Q4



Response of Output Gap



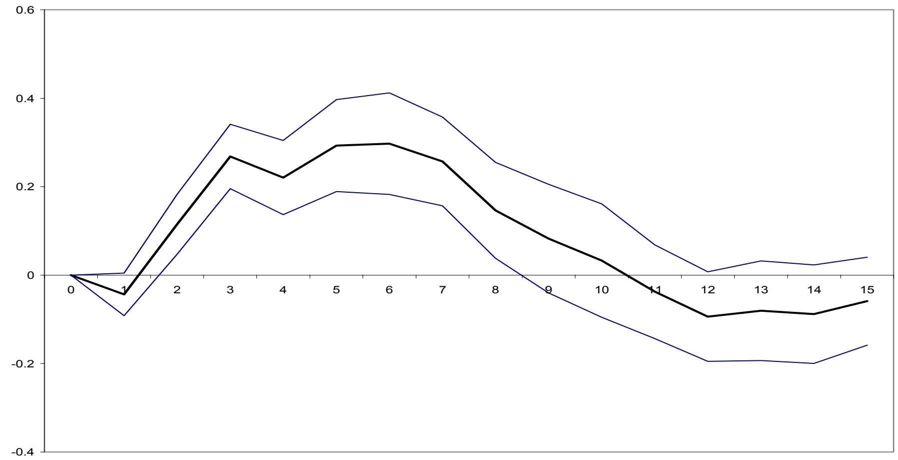
Response of Inflation



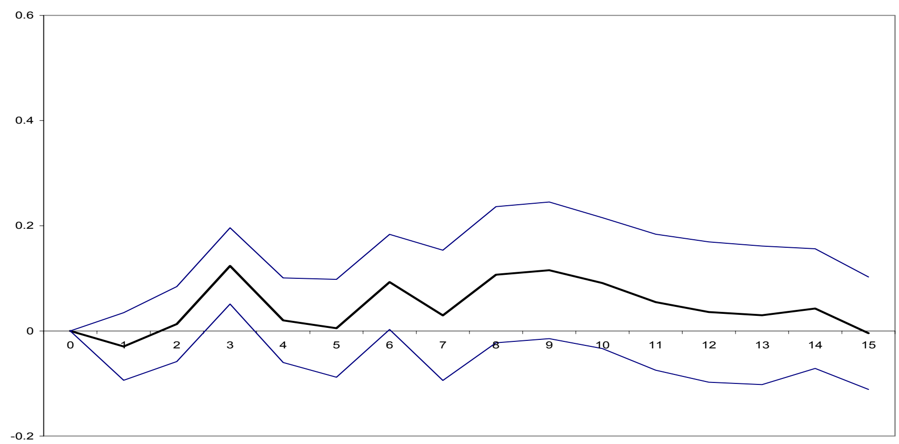
Response of FFR

FIGURE 15

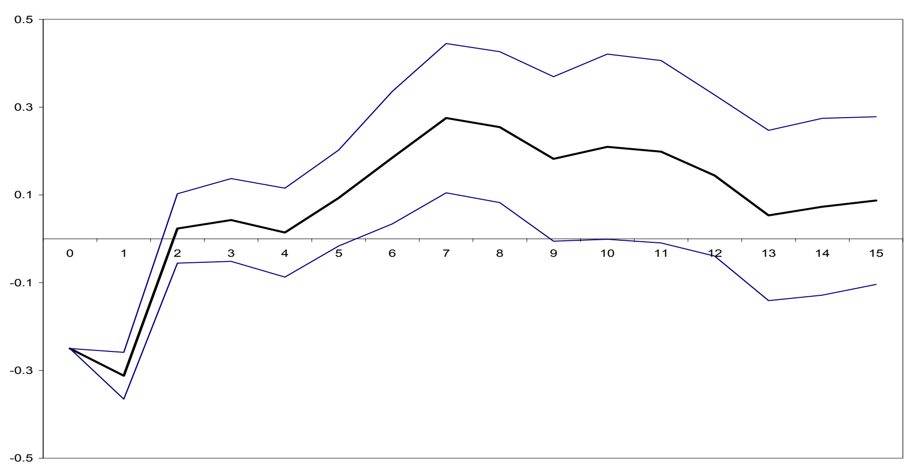
25-Basis Point Decline in Fed Funds Rate in Q2
Quarterly Dependence. Output-Gap System 1966:Q1 to 2002:Q4



Response of Output Gap



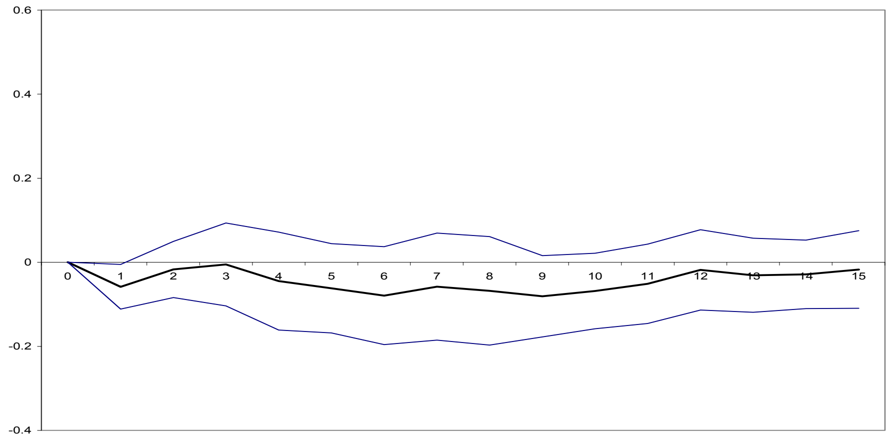
Response of Inflation



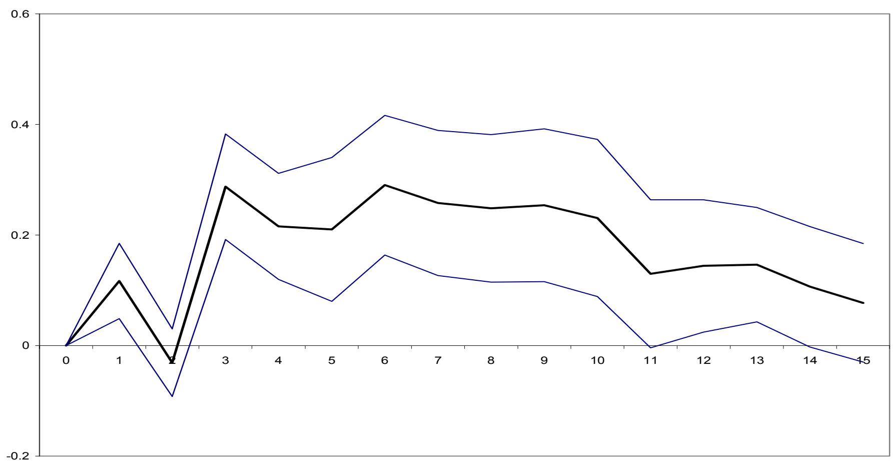
Response of FFR

FIGURE 16

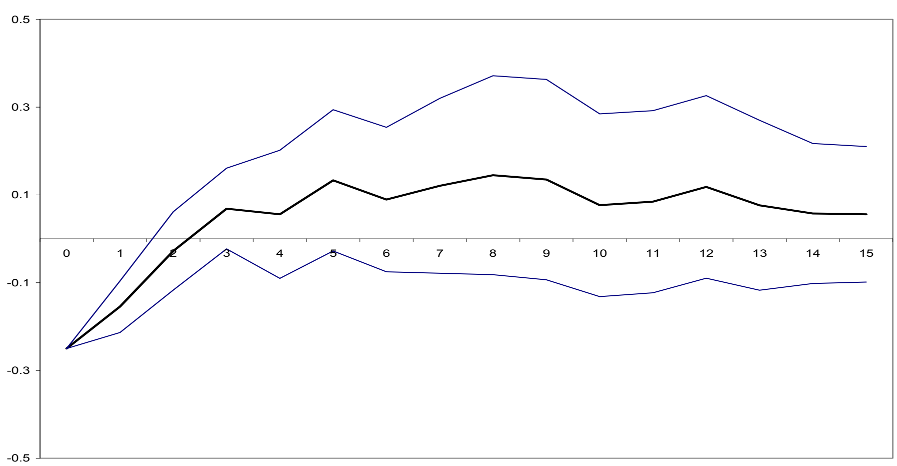
25-Basis Point Decline in Fed Funds Rate in Q3
Quarterly Dependence. Output-Gap System 1966:Q1 to 2002:Q4



Response of Output Gap



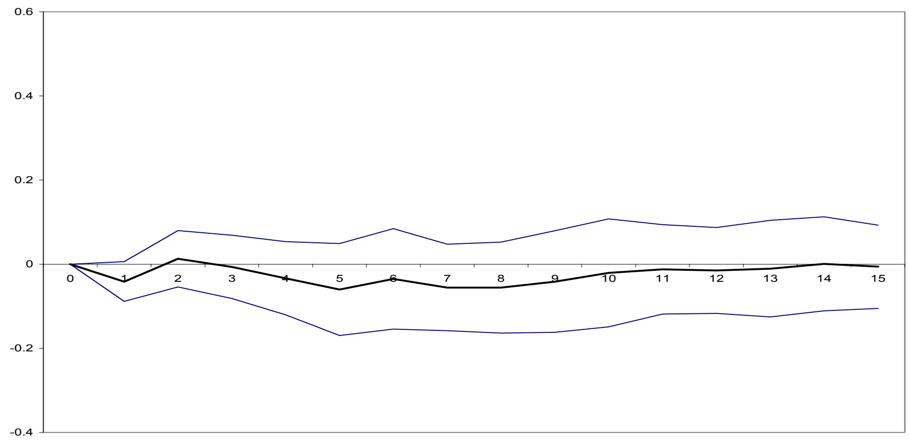
Response of Inflation



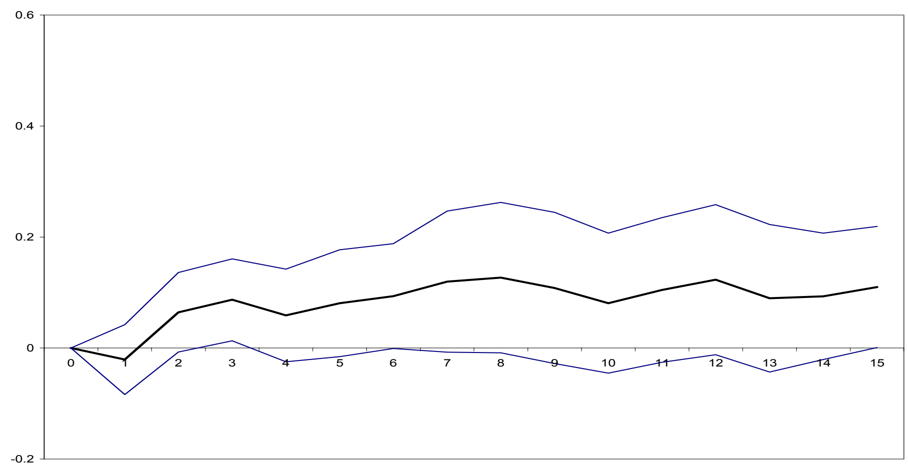
Response of FFR

FIGURE 17

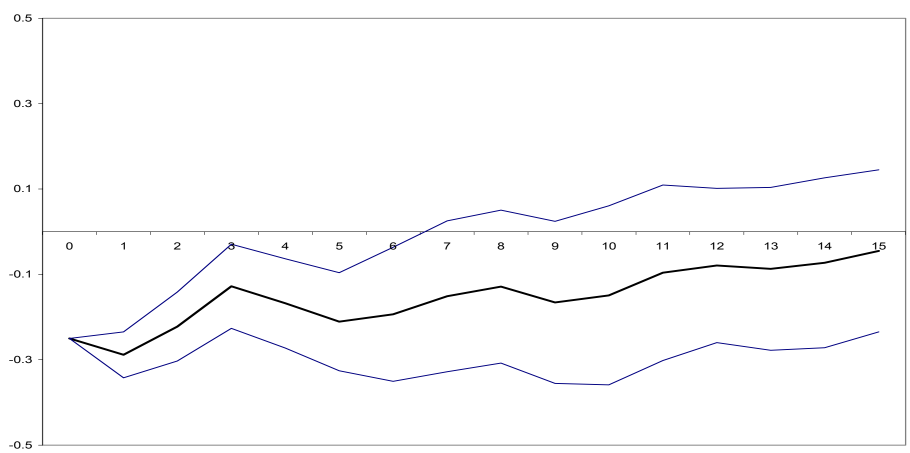
25-Basis Point Decline in Fed Funds Rate in Q4
Quarterly Dependence. Output-Gap System 1966:Q1 to 2002:Q4



Response of Output Gap



Response of Inflation



Response of FFR

Table 1. Differences in Impulse-Responses Across Quarters
p-values for D-Statistic

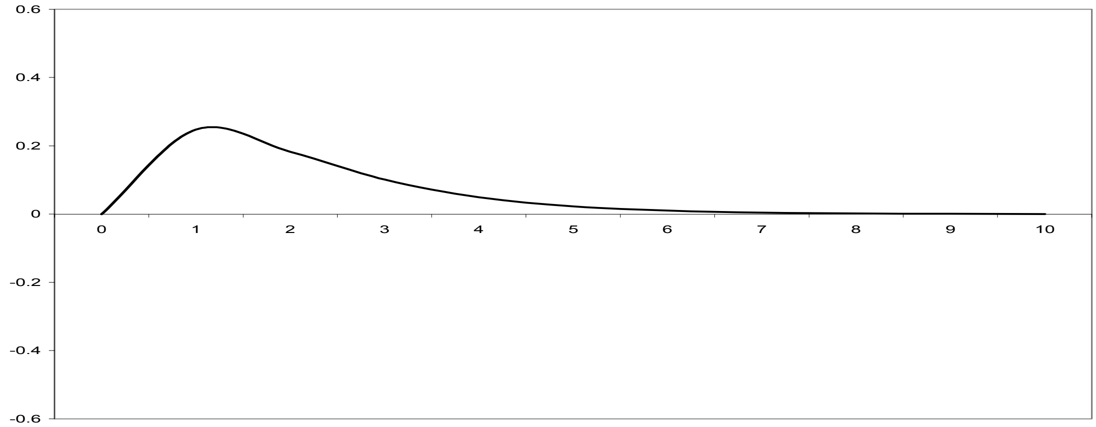
Variable	Quarter			
	First	Second	Third	Fourth
GDP	0.01	0.00	0.06	0.40
GDP Deflator	0.40	0.70	0.10	0.55
Fed Funds Rate	0.20	0.01	0.15	0.60

Table 2. Parameter Calibration

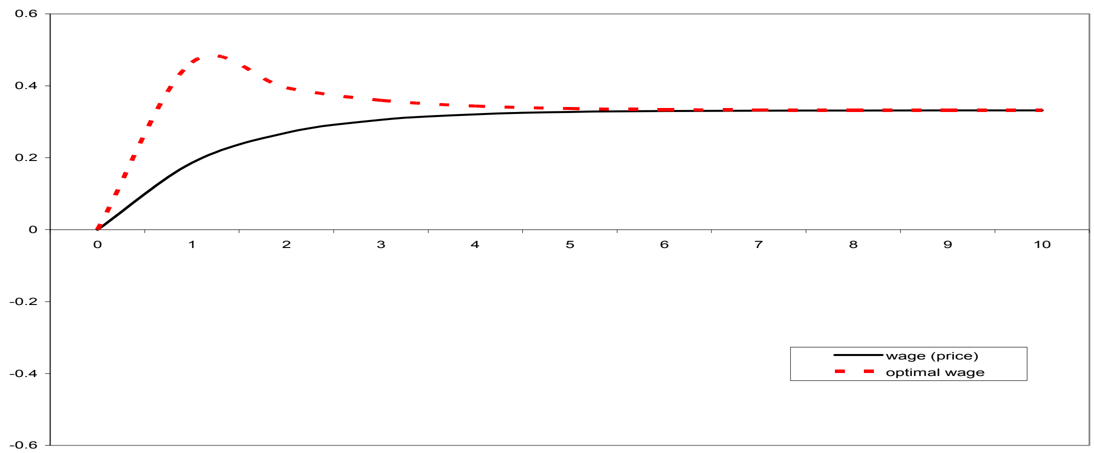
Parameters	Uniform Staggering	Uneven Staggering
β	0.95	0.95
δ_p	0.42	0.42
δ_y	0.11	0.11
ξ	0.84	0.84
ψ	-0.10	-0.10
η	0.30	0.30
α_1	0.60	0.70
α_2	0.60	0.90
α_3	0.60	0.50
α_4	0.60	0.40

FIGURE 18

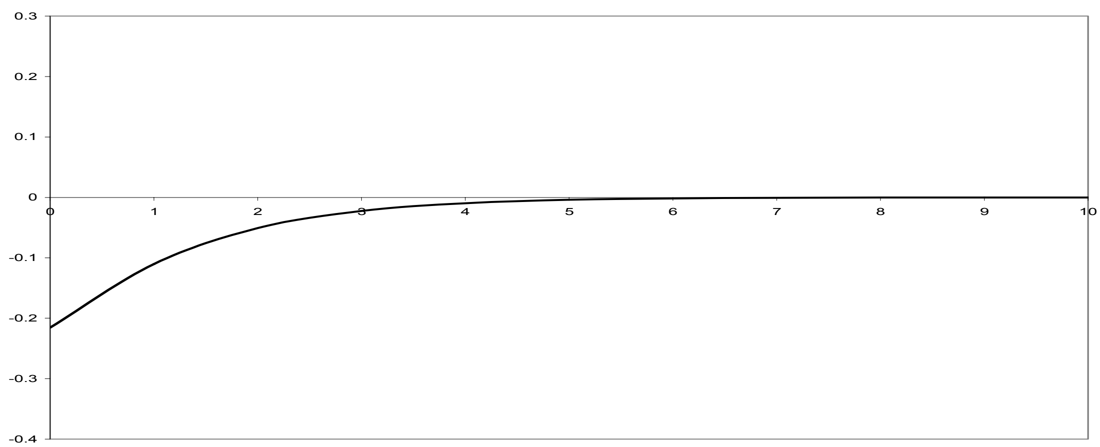
Simulations: Decline in Fed Funds Rate
No Quarterly Dependence.



Response of Output



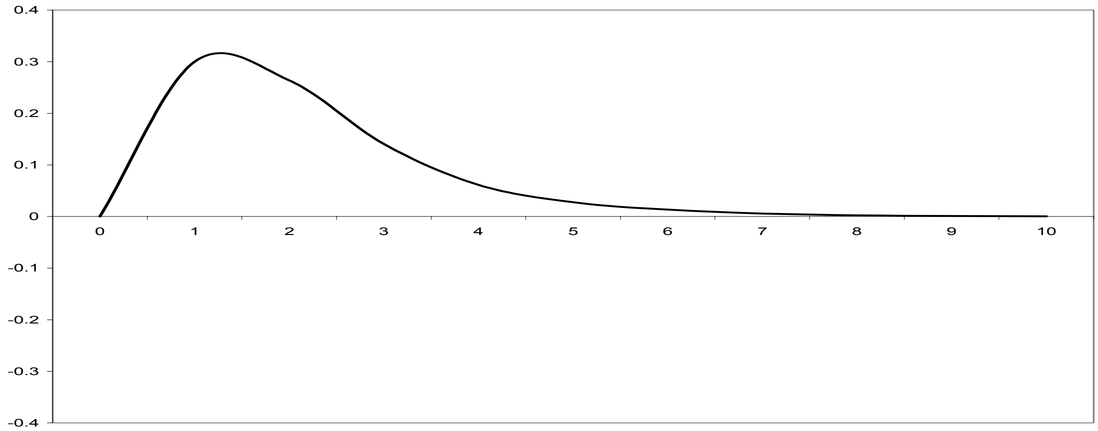
Response of Wage/Price and Optimal Wage



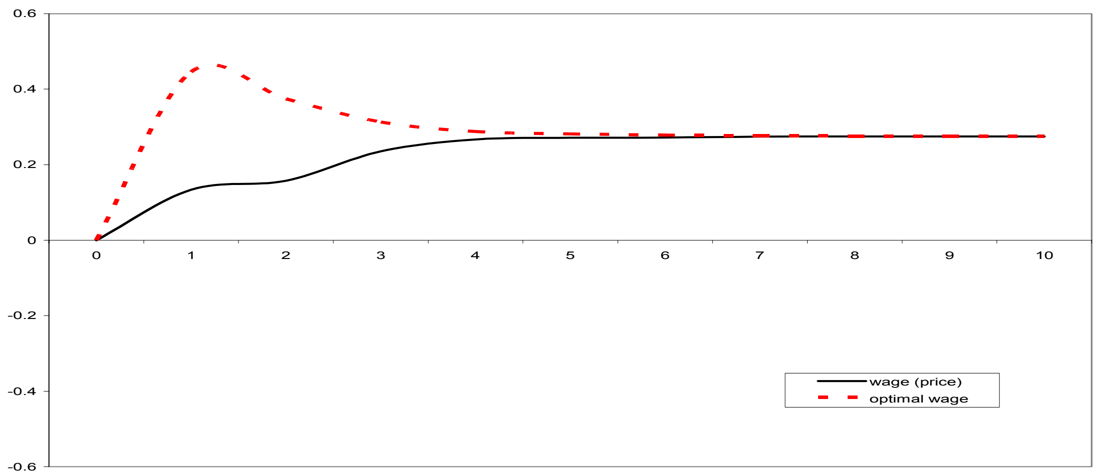
Response of Nominal Interest Rate

FIGURE 19

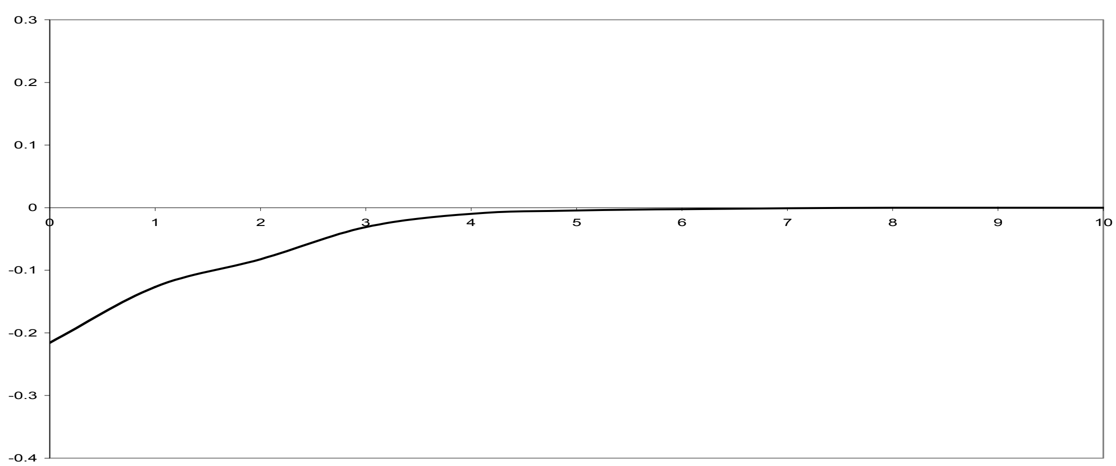
Simulations: Decline in Fed Funds Rate in Q1
Quarterly Dependence.



Response of Output



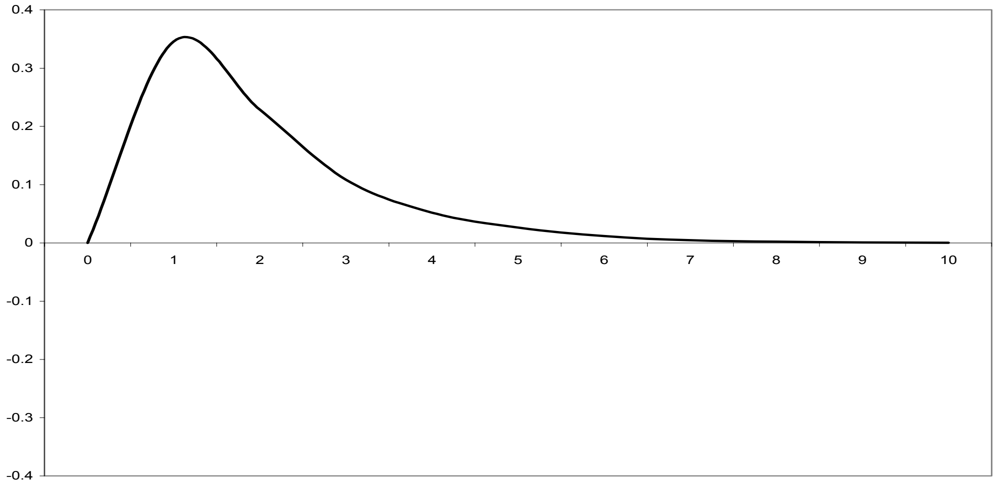
Response of Wage/Price and Optimal Wage



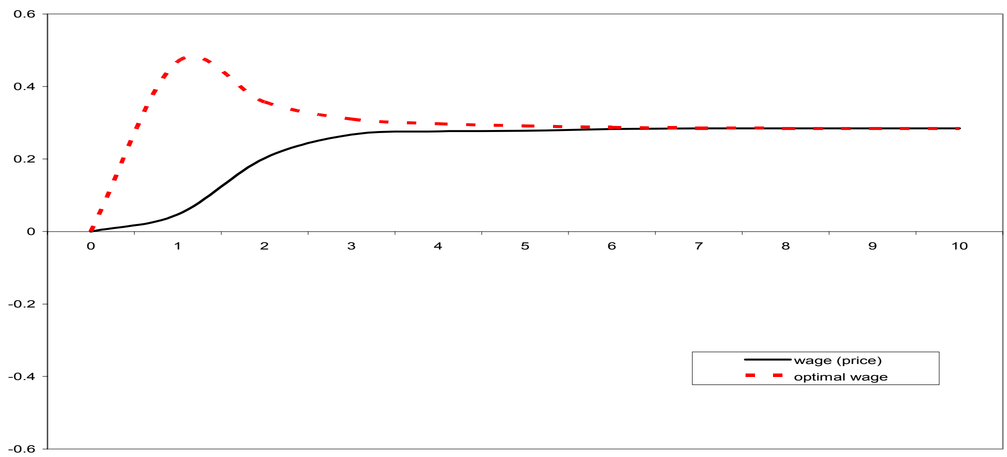
Response of Nominal Interest Rate

FIGURE 20

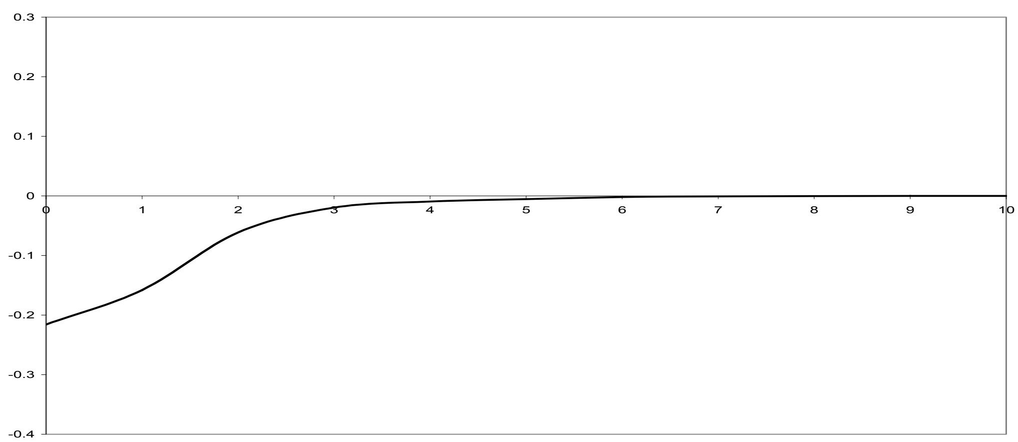
Simulations: Decline in Fed Funds Rate in Q2
Quarterly Dependence.



Response of Output



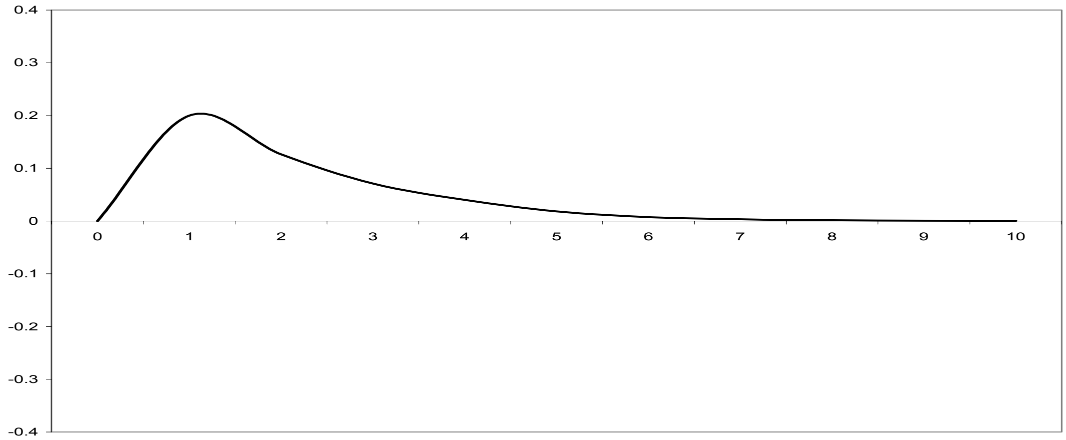
Response of Wage/Price and Optimal Wage



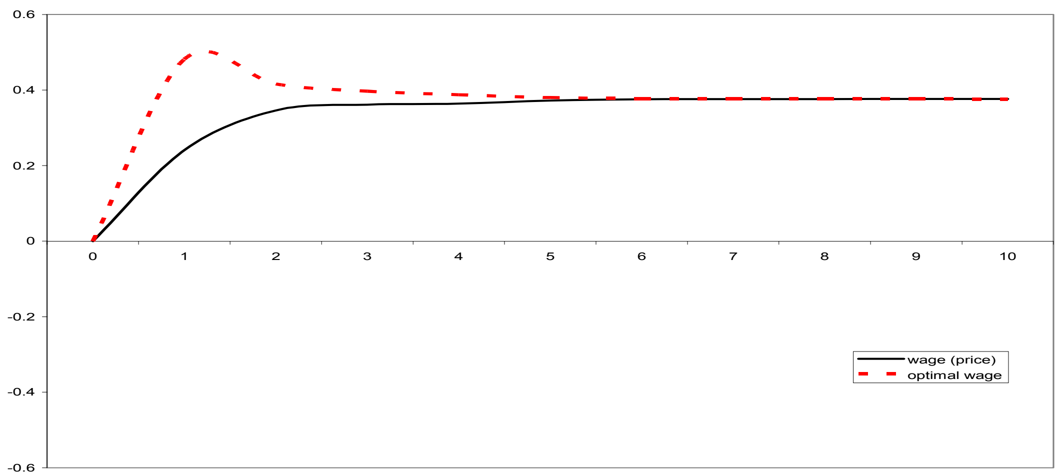
Response of Nominal Interest Rate

FIGURE 21

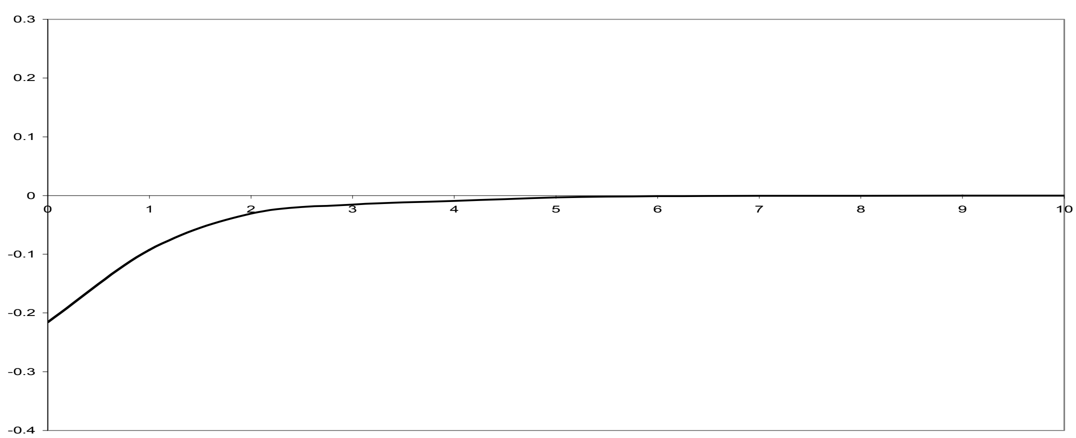
Simulations: Decline in Fed Funds Rate in Q3
Quarterly Dependence.



Response of Output



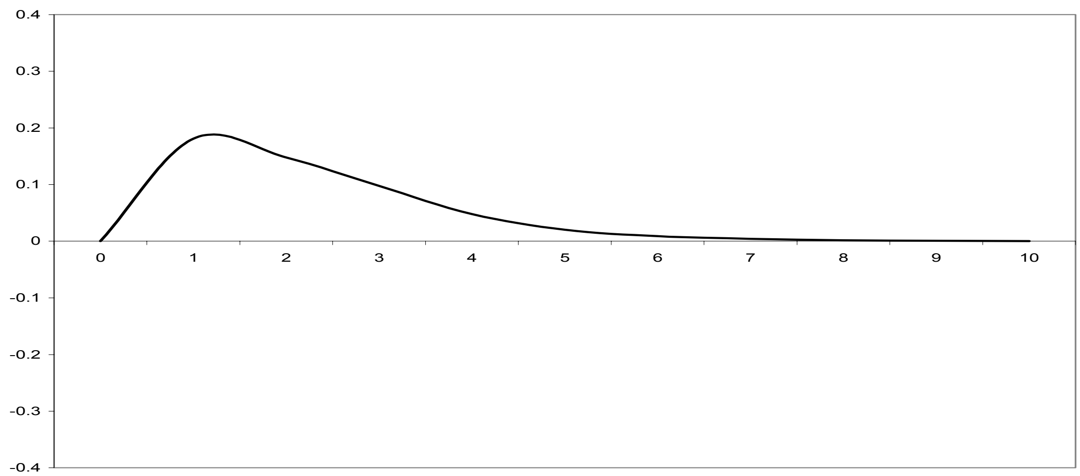
Response of Wage/Price and Optimal Wage



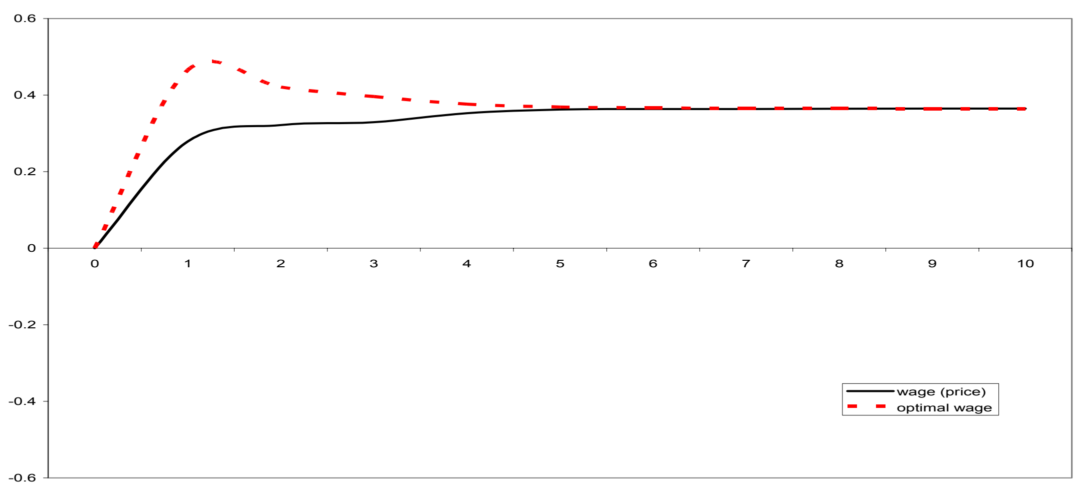
Response of Nominal Interest Rate

FIGURE 22

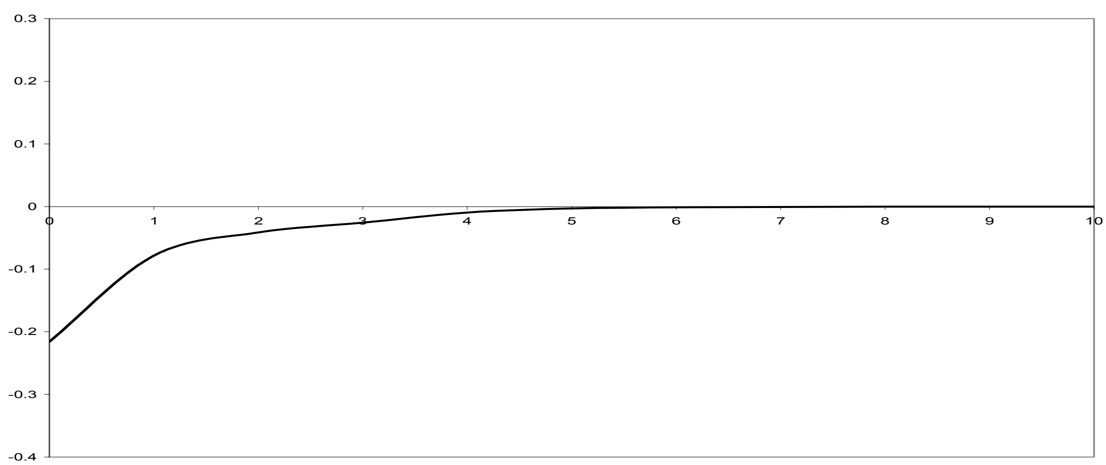
Simulations: Decline in Fed Funds Rate in Q4
Quarterly Dependence.



Response of Output



Response of Wage/Price and Optimal Wage



Response of Nominal Interest Rate