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Inflation Measure, Taylor Rules, and the Greenspan-Bernanke Years

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Recent research has highlighted several aspects of monetary policy under Chairman Alan Greenspan, noting that the Federal Reserve was forward looking, smoothed interest rates, and focused on core inflation.[1] Some analysts have estimated Taylor rules that incorporate these salient features of monetary policy, and have shown that monetary policy actions taken by the Federal Reserve in the Greenspan era can broadly be explained by these estimated Taylor rules. Using a core measure of consumer price inflation (CPI), Blinder and Reis (2005) estimate a Taylor rule over 1987:1–2005:1, showing that the estimated policy rule tracks actual policy actions fairly well. Using Greenbook forecasts of core CPI inflation, Mehra and Minton (2007) estimate a forecast-based Taylor rule that shows this estimated policy rule also fits the data over 1987:1–2000:4.[2] More recently however, Taylor (2007, 2009) has argued that monetary policy was "too loose" during most of the period from 2002–2006, in the sense that the actual federal funds rate was too low relative to the level simulated by a smoothed version of the original Taylor rule.[3] In this simulation exercise, Taylor (2007) assumes response coefficients

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[1] Blinder and Reis (2005) have hailed Chairman Greenspan's focus on core, rather than headline, inflation as a "Greenspan innovation." The measure of core inflation used excludes food and energy prices.

[2] The sample period used in Mehra and Minton (2007) ends in 2000, given the five-year lag in the release of the Greenbook forecasts to the public.

[3] The original Taylor rule relates the federal funds rate target to two economic variables—lagged inflation and the output gap, with the actual federal funds rate completely adjusting to the

of 1.5 and .5 on inflation and the output gap, as in the original Taylor rule, but instead uses headline CPI as a measure of inflation.[4]

This article highlights another aspect of monetary policy in the Greenspan era: The measure of inflation used in monetary policy deliberations has also been refined over time. This can be seen in the semiannual monetary policy reports to Congress (Humphrey-Hawkins reports), where inflation forecasts by the members of the Federal Open Market Committee (FOMC) have been presented using different measures of inflation over time. Thus, through July 1988, inflation forecasts used the implicit deflator of the gross national product, thereafter switching to the CPI. In February 2000, the CPI was replaced by the personal consumption expenditures (PCE) deflator measure of inflation and from July 2004 onward inflation forecasts employed the core PCE deflator that excludes food and energy prices.

Though these different measures of inflation may move together in the long run, over short periods these inflation measures may behave differently because of factors such as energy prices and changes in coverage and definitions. As a result, the Fed's inflation target may vary depending on the measure of inflation used, thereby affecting the desired setting of the federal funds rate.[5] Previous empirical work has not paid much attention to this issue, as most analysts estimate Taylor rules under the assumption that the measure of inflation used in policy deliberations did not change during the Greenspan years.[6]

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target each period as shown below (Taylor 1993):

$$\begin{aligned} FR_t &= rr^* + \alpha_\pi \left(\pi_{t-1} - \pi^*\right) + \alpha_y \left(y_t - y_t^*\right)_{t-1} \\ FR_t &= 2.0 + 1.5 \left(\pi_{t-1} - 2.0\right) + .5 \left(y_t - y_t^*\right)_{t-1}, \end{aligned}$$

where rr^* is the real interest rate (assumed to be 2 percent), π is actual inflation, π^* is the Fed's inflation target (assumed to be 2 percent), $\left(y_t - y_t^*\right)$ is the output gap, α_π is the inflation response coefficient (assumed to be 1.5), and α_y is the output response coefficient (assumed to be .5). Inflation in the original Taylor rule was measured by the behavior of the gross domestic product (GDP) deflator, and the output gap is the deviation of the log of real output from a linear trend. According to the original Taylor rule, the Federal Reserve is backward looking, focused on headline inflation, and follows a "non-inertial" policy rule.

[4] Using the policy response coefficients from the original Taylor rule and headline CPI as a measure of inflation, Poole (2007) also shows that the actual federal funds rate is too low relative to the level prescribed for most of the period from 2000:1–2006:4 (see Figure 1 in Poole [2007]). Poole, however, does not conclude that policy was too easy because he alludes to the change in the measure of inflation used in monetary policy deliberations during this subperiod.

[5] Kohn (2007) has highlighted these considerations.

[6] While Blinder and Reis (2005) and Mehra and Minton (2007) estimate Taylor rules using a core measure of CPI inflation, the measure of inflation used in Taylor (2007, 2009) is headline CPI and the one used in Smith and Taylor (2007) is the implicit deflator for GDP. An exception is the paper by Orphanides and Wieland (2008), in which a forecast-based Taylor rule is estimated using the semiannual Humphrey-Hawkins inflation forecasts. More recently, Dokko et al. (2009) and Bernanke (2010) have highlighted the issue of the measurement of inflation used in monetary policy deliberations.

This article re-examines the issue of whether monetary policy actions taken during the Greenspan years can be described by a stable Taylor rule. It considers two Taylor rules that differ with respect to the measure of inflation used in implementing monetary policy. According to both these rules, the Greenspan Fed was forward looking, smoothed interest rates, and linked the federal funds rate target to expected inflation and the unemployment gap. However, according to one Taylor rule, the Federal Reserve used headline CPI inflation, and, according to the other, it used core CPI until 2000 and core PCE thereafter. The later specification departs from the usual assumption that a Taylor rule has to be estimated using a single measure of inflation.[7] Both the policy rules employ real-time data on economic fundamentals such as the pertinent inflation measure and the unemployment gap. As noted by Orphanides (2001, 2002), in evaluating historical monetary policy actions using estimated Taylor rules, the use of ex post revised, as opposed to real-time, data on economic variables can give misleading inferences about the stance of monetary policy.

A Taylor rule incorporating the above-noted features is shown below in (1.3):

$$FR_t^* = \alpha_0 + \alpha_\pi E_t \pi_{t+j}^c + \alpha_u \left(ur_t - ur_t^* \right), \tag{1.1}$$

$$FR_t = \rho FR_{t-1} + (1-\rho)\, FR_t^* + \nu_t, \tag{1.2}$$

$$FR_t \quad = \quad \rho FR_{t-1} + (1-\rho) \left\{ \alpha_0 + \alpha_\pi E_t \pi_{t+j}^c + \alpha_u \left(ur_t - ur_t^* \right) \right\} + \nu_t, \tag{1.3}$$

where FR_t is the actual federal funds rate, FR_t^* is the federal funds rate target, $E_t \pi_{t+j}^c$ is the expectation of the j-period-ahead core inflation rate made at time t conditional on period $t-1$ dated information, ur is the actual unemployment rate, ur^* is the non-accelerating inflation unemployment rate (NAIRU), and ν_t is the disturbance term. Thus, the term $\left(ur_t - ur_t^* \right)$ is the current unemployment gap. Equation (1.1) relates the federal funds rate target to two economic fundamentals, expected inflation and the current unemployment gap. Hereafter, the funds rate target is called the policy rate. The coefficients α_π and α_u measure the long-term responses of the funds rate target to expected inflation and the unemployment gap; the inflation response coefficient is assumed to

[7] The estimation of a Taylor rule using an inflation series that employs two or more measures of inflation may mean that the intercept term in the estimated Taylor rule is no longer a constant. This may happen if different measures of inflation exhibit different trend behaviors during the course of the estimation period and, hence, the Fed's inflation target expressed in these different inflation measures is no longer similar in magnitude.

be positive and the unemployment gap response coefficient is assumed to be negative, indicating that the Federal Reserve raises its funds rate target if it expects inflation to rise and/or the unemployment gap to fall. Equation (1.2) is the standard partial adjustment equation, which expresses the current funds rate as a weighted average of the current funds rate target, FR_t^*, and the last quarter's actual value, FR_{t-1}. If the actual funds rate adjusts to its target within each period, then ρ equals zero, suggesting that the Federal Reserve does not smooth interest rates. Equation (1.2) also includes a disturbance term, indicating that in the short run the actual funds rate may deviate from the value implied by economic determinants specified in the policy rule. If we substitute (1.1) into (1.2), we get (1.3)—a forward-looking "inertial" Taylor rule.

As in Clarida, Gali, and Gertler (2000), the Taylor rules are estimated assuming rational expectations and using instrumental variables over 1987:1–2004:4; this sample period spans most of the Greenspan era.[8] The key feature of the estimation procedure used here is that the instrument set includes, among other variables, Greenbook inflation forecasts based on different inflation measures. This strategy differs from the one used in Boivin (2006) and Mehra and Minton (2007), where forward-looking Taylor rules are estimated directly using Greenbook forecasts. Given the five-year lag in the release of Greenbook forecasts to the public, the current strategy enables one to estimate the Taylor rules over most of the Greenspan era (1987:1–2004:4) and then examine their predictive content for the longer sample period (1987:1–2006:4) that includes the Bernanke years.[9] We end the sample in 2006 in order to compare results in previous research that indicate monetary policy was too loose over 2002–2006.

The empirical work presented here suggests several observations. First, a Taylor rule that is estimated using a time-varying measure of core inflation

[8] There is considerable evidence that the policy rule followed by the Greenspan Fed differed from the one followed by the Volcker Fed in one important way. In its attempts to build credibility, the Volcker Fed responded strongly to long-term inflationary expectations imbedded in long bond yields, in addition to responding to inflation and unemployment, the two fundamental variables suggested by a Taylor rule (Mehra 2001). The long bond rate is generally not significant if the Taylor rule is estimated using data from the Greenspan era because the Greenspan Fed had by then achieved credibility. For this reason we estimate the Taylor rule using observations only from the Greenspan era. This strategy is also consistent with the observation that in criticizing the Greenspan Fed, Taylor (2007) uses a policy rule that includes only inflation and unemployment (output) gap variables.

[9] We, however, do compare the robustness of our results to this alternative method of estimating the Taylor rule using Greenbook inflation forecasts. Although estimates of policy response coefficients differ, the estimates yield qualitatively similar conclusions about the relevance of the inflation measure. In particular, the Taylor rule that is estimated using Greenbook forecasts of core CPI until 2000 and core PCE thereafter tracks actual policy well over 2000:1–2006:4 and passes the test of parameter stability. That is not the case if the Taylor rule is estimated using Greenbook forecasts of headline CPI inflation. Furthermore, as measured by the root mean squared error criterion, the Taylor rule with Greenbook forecasts of the time-varying inflation measure fits the data better than the Taylor rule with Greenbook forecasts of headline CPI.

(CPI until 2000 and PCE thereafter) yields reasonable estimates of inflation and unemployment gap response coefficients. The estimated inflation response coefficient, α_{π}, is positive and way above unity, suggesting that the Greenspan Fed responded strongly to expected inflation. The estimated unemployment gap response coefficient, α_u, is negative and statistically significant, suggesting that the Federal Reserve also responded to slack. The Chow test of parameter stability does not indicate a shift in the estimated parameters around 2000 when the Federal Reserve switched from CPI to PCE.[10] Also, the estimated Taylor rule tracks the actual path of the federal funds rate fairly well, especially over the period from 2002–2006.

In contrast, a Taylor rule that is estimated using headline CPI inflation does not provide reasonable estimates of policy response coefficients and depicts parameter instability over 1988:1–2004:4. The estimated Taylor rule based on headline CPI inflation is consistent with the actual funds rate being too low relative to the level prescribed by the estimated Taylor rule over 2002–2006, as in Smith and Taylor (2007) and Taylor (2007). These results indicate that the choice of the measure of inflation used in estimated Taylor rules is not innocuous. Furthermore, one should employ real-time information for evaluating historical monetary policy actions.

Second, during most of the period from 2001–2006, inflation measured by headline CPI was higher than what would be indicated by core PCE data, reflecting in part the effects of the rise in oil prices on headline inflation. The tests of parameter stability here indicate that the Greenspan Fed did not adjust the federal funds rate target in response to increases in the headline measure of CPI inflation.[11] The lack of policy response to increases in headline CPI inflation reflected the Greenspan Fed's belief that oil price increases were transitory[12] and that core inflation is a better gauge of the underlying trend inflation.[13]

Third, the core measure of PCE inflation has been substantially revised over the years. In particular, real-time estimates of core PCE inflation over

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[10] Although the core PCE index was given prominence in Humphrey-Hawkins forecasts in July 2004, the hypothesis here that the Greenspan Fed in fact paid attention to core measures of inflation implies that the FOMC started paying attention to core PCE much earlier.

[11] Several analysts and policymakers have noted that the Greenspan Fed's policy of focusing on core inflation continued through the Bernanke years. See, for example, Kohn (2009) and Bernanke (2010).

[12] During this subperiod most other economists also thought oil price increases were transitory and hence did not expect the rise in oil prices to lead to persistent increases in headline inflation. For example, despite the actual increase in headline CPI inflation, the Survey of Professional Forecasters forecasts of headline CPI inflation did not increase appreciably over 2003:1–2006:4. See Dokko et al. (2009) for additional evidence on this issue.

[13] This belief is consistent with the empirical evidence documented by several analysts that, for the period since the early 1980s, it is core rather than headline inflation that better approximates the trend component of inflation. Some of that empirical evidence is reviewed in Mishkin (2007) and Kiley (2008) and updated in Mehra and Reilly (2009).

2002:1–2005:4 are substantially lower than those indicated by ex post revised data (vintage 2009). The counterfactual simulations of the federal funds rate generated using the ex post revised data do suggest that deviations of the policy rule are somewhat larger than those generated using the real-time data. However, it would be misleading to conclude from such evidence that the Greenspan Fed had followed an easier stance on monetary policy.

Our results complement the recent work of Orphanides and Wieland (2008), who argue that policy actions taken over 1988–2007 have been consistent with a stable Taylor rule and that policy was not too loose over 2001–2007. They, however, estimate a forecast-based Taylor rule using publicly available forecasts of inflation and unemployment contained in semiannual Humphrey-Hawkins reports. As indicated before, the Humphrey-Hawkins inflation forecasts used CPI until 1999, switching thereafter to the PCE measure. The evidence in this article implies that a forward-looking Taylor rule estimated using actual real-time inflation and unemployment data yields identical results, in particular the conclusion that policy actions are consistent with a stable Taylor rule, provided we allow for the change in the measure of inflation used in monetary policy deliberations.[14]

The rest of the paper is organized as follows. Section 1 discusses the empirical methodology and reviews the data on the behavior of different measures of inflation during the Greenspan era. Section 2 presents empirical results, reproducing the evidence in Taylor (2007, 2009) that the Greenspan Fed set a funds rate low relative to the Taylor rule. We show that the result in Taylor disappears if one uses the time-varying measure of inflation employed by the FOMC. Section 3 concludes.

1. EMPIRICAL METHODOLOGY

Estimation of the Forward-Looking Inertial Taylor Rule

The objective of this article is to investigate whether monetary policy actions taken by the Federal Reserve under Chairman Greenspan can be summarized by a Taylor rule according to which the Federal Reserve was forward looking, focused on core inflation, smoothed interest rates, and refined the measure of inflation used in monetary policy deliberations. We model the forward-looking nature of the policy rule by relating the current value of the funds rate

[14] Using somewhat different approaches, Dokko et al. (2009) and Bernanke (2010) also show that actual policy is much closer to the one prescribed by the original Taylor rule if the measure of inflation used in the policy rule is the one employed by the FOMC in monetary policy deliberations and if real-time data are used. Bernanke (2010) generates the predictions of the policy rate using the Greenbook inflation forecasts until 2004 and the Humphrey-Hawkins forecasts thereafter. Dokko et al. (2009) generate the predictions of the policy rate employing real-time estimates of core PCE inflation.

target to the expected average annual inflation rate and the contemporaneous unemployment gap. The policy rule incorporating these features is reproduced below in equation (2.3):

$$FR_t = \rho FR_{t-1} + (1-\rho)\left\{\alpha_0 + \alpha_\pi E_t \bar{\pi}^c_{\bar{4}} + \alpha_u \left(ur_t - ur^*_t\right)\right\} + \nu_t, \quad (2.3)$$

where the expected average annual inflation rate, $E_t \bar{\pi}^c_{\bar{4}}$, is measured by the average of one-through-four-quarter-ahead expected values of core inflation made at time t, and other variables are defined as before.[15]

The estimation of the policy rule (2.3) raises several issues. The first issue relates to how we measure expected inflation and the unemployment gap. The second issue relates to the nature of data used in estimation, namely, whether it is the real-time or final revised data. As indicated earlier, the use of revised as opposed to real-time data may affect estimates of policy coefficients and may provide a misleading historical analysis of policy actions. The third issue is an econometric one, arising as a result of the potential presence of serial correlation in the error term ν_t. Rudebusch (2002, 2006) points out that the Federal Reserve may respond to other economic factors besides expected inflation and the unemployment gap and, hence, a Taylor rule estimated while omitting those other factors is likely to have a serially correlated error term. The presence of serial correlation in the disturbance term, if ignored, may spuriously indicate that the Federal Reserve is smoothing interest rates.

To understand how a serially correlated disturbance term may mistakenly indicate the presence of partial adjustment, note first that if the funds rate does partially adjust to the policy rate as shown in (1.2) and the disturbance term has no serial correlation, then the reduced-form policy rule in (1.3) or (2.3) has the lagged funds rate as one of the explanatory variables. Hence, the empirical finding of a significant coefficient on the lagged funds rate in the estimated policy rule may be interpreted as indicating the presence of interest rate smoothing. But now assume that there is no partial adjustment, $\rho = 0$ in (2.3), but instead the disturbance term is serially correlated as shown below in (3.1):

$$\nu_t = s\nu_{t-1} + \varepsilon_t, \quad (3.1)$$

[15] In particular, the four-quarter average of expected inflation rates is defined as $\bar{\pi}^c_{t,\bar{4}} = \left(\pi^c_{t,1} + \pi^c_{t,2} + \pi^c_{t,3} + \pi^c_{t,4}\right)/4$, where $\pi^c_{t,j}\, j = 1, 2, 3, 4$ is the j-quarter-ahead expected value of core inflation made at time t.

$$FR_t = sFR_{t-1} + \left\{\alpha_0 + \alpha_\pi E_t \bar{\pi}_4^c + \alpha_u \left(ur_t - ur_t^*\right)\right\} \\ -s\left\{\alpha_0 + \alpha_\pi E_{t-1}\bar{\pi}_4^c + \alpha_u \left(ur_{t-1} - ur_{t-1}^*\right)\right\} + \varepsilon_t. \quad (3.2)$$

If we substitute (3.1) into (2.3), it can be easily shown that we get the reduced-form policy rule (3.2) in which, among other variables, the lagged funds rate also enters the policy rule. Hence, the empirical finding of a significant coefficient on the lagged funds rate in the estimated policy rule may be interpreted as arising as a result of interest rate smoothing when in fact it is not present. In view of these considerations, the policy rule here is estimated allowing for the presence of both interest rate smoothing and serial correlation, namely, we allow both partial adjustment and a serially correlated disturbance term. It can be easily shown that the policy rule incorporating both partial adjustment and serial correlation can be expressed

$$FR_t = \alpha_0(1-s)(1-\rho) + (s+\rho)FR_{t-1} \\ + (1-\rho)\left\{\alpha_\pi E_t\bar{\pi}_4^c + \alpha_u\left(ur_t - ur_t^*\right)\right\} \\ -s\left\{(1-\rho)\alpha_\pi E_{t-1}\bar{\pi}_4^c + (1-\rho)\alpha_u\left(ur_{t-1} - ur_{t-1}^*\right)\right\} \\ -s\rho FR_{t-2} + \varepsilon_t. \quad (4)$$

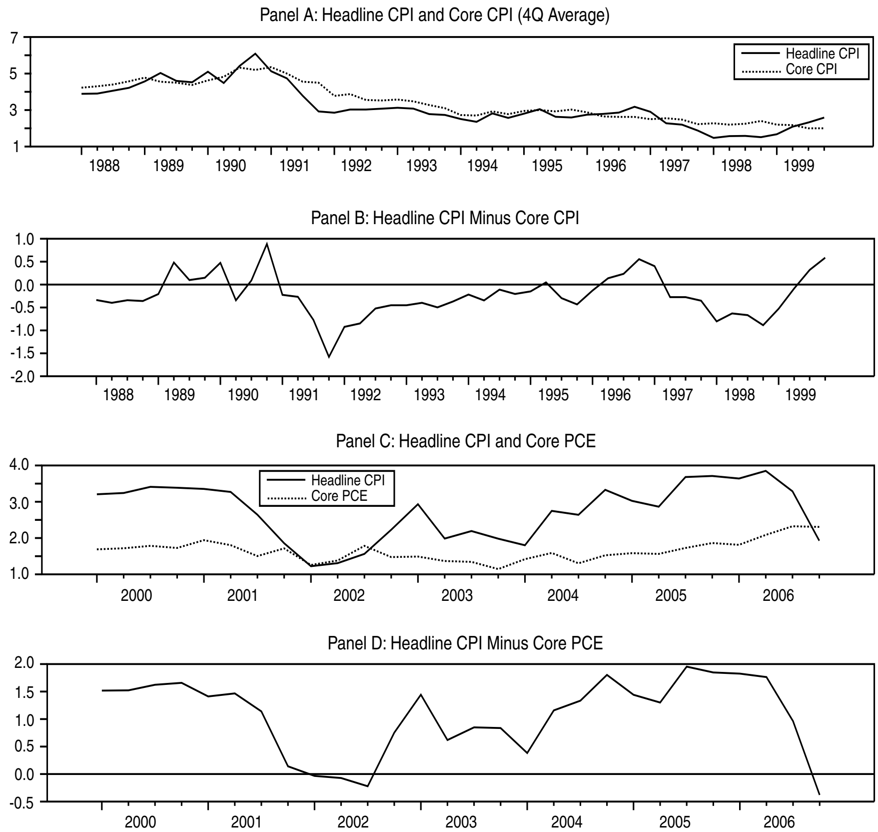
Note, if there is no serial correlation ($s = 0$ in [4]), we get the reduced-form policy rule shown in (2.3), and if there is no partial adjustment ($\rho = 0$ in [4]), we get the policy rule shown in (3.2). Of course, if both s and ρ are not zero, we have a policy rule with both partial adjustment and serial correlation.[16]

In previous research, a forward-looking policy rule such as the one given in (2.3) has often been estimated assuming rational expectations and using a generalized method of moments procedure (Clarida, Gali, and Gertler 2000). We follow this literature and estimate the policy rule assuming rational expectations, namely, we substitute actual future core inflation for the expected inflation term and use an instrumental variables procedure to estimate policy coefficients. Given the evidence that the Greenbook forecasts are most appropriate in capturing policymakers' real-time assessment of future inflation developments, we include the Greenbook forecasts in the instruments.[17] In

[16] Estimating the policy rule allowing for the presence of serial correlation produces more robust estimates of policy parameters including the partial adjustment coefficient. Moreover, the policy rule is estimated using the quasi-differenced data, as can be seen in equation (3.2). This quasi-differencing of data minimizes the spurious regression phenomenon noted in Granger and Newbold (1974).

[17] Romer and Romer (2000) have shown that the Federal Reserve has an informational advantage over the private sector, producing relatively more accurate forecasts of inflation than does the private sector. Bernanke and Boivin (2003) argue that one needs a large set of conditional information to properly model monetary policy. In that respect, the Greenbook forecasts include

Figure 1 Actual Inflation (Real-Time)

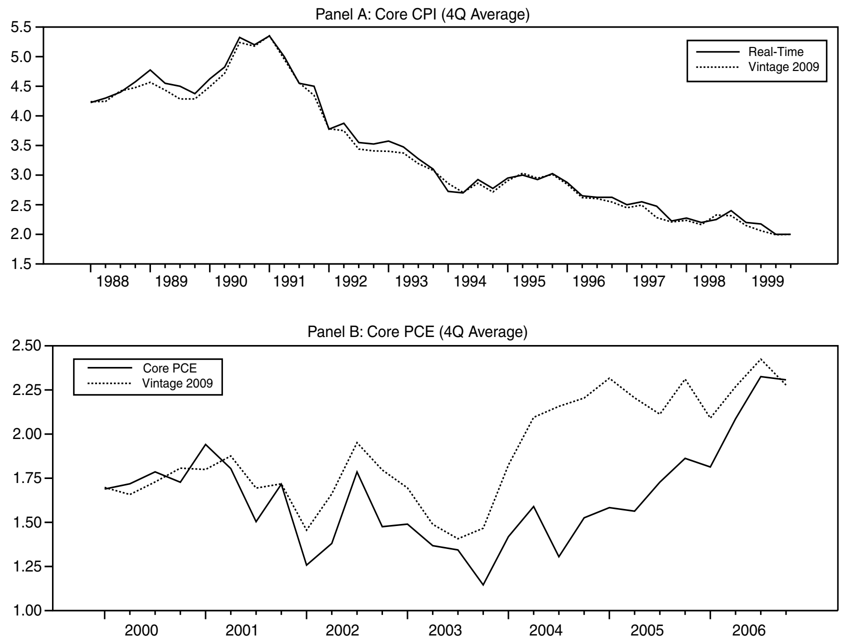


addition, we estimate the policy rule allowing for the presence of both interest rate smoothing and serial correlation as in (4) and use the nonlinear instrumental variables procedure. The instruments used are the three lagged values of Greenbook inflation forecasts, the federal funds rate, levels of the unemployment gap, and the spread between the 10-year Treasury bond yield and the federal funds rate. As indicated earlier, the policy rule is estimated over 1988:1–2004:4, given the five-year lag in the release of the Greenbook forecasts to the public.[18]

real-time information from a wide range of sources, including the Board staff's "judgment," not otherwise directly measurable.

[18] The estimation period begins in 1988:1 because the instrument set includes the lagged values of economic variables. As a check on the adequacy of the instruments variables procedure, we ran the first-stage regressions for the endogenous variables (expected inflation and the

Figure 2 Real-Time Versus Vintage 2009 Core Inflation



Data

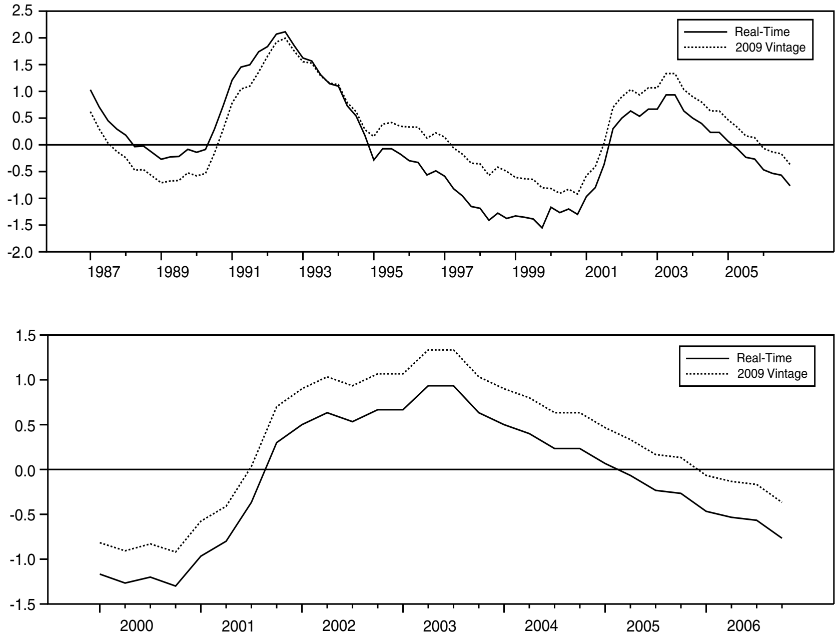
We estimate the policy rule in (4) using real-time data on core inflation and the unemployment gap. The data on core inflation came from the real-time data set maintained at the Philadelphia Fed.[19] The data on real-time estimates of the NAIRU were those prepared by the Congressional Budget Office (CBO).[20] The Greenbook forecasts of core inflation used in the instrument list are those prepared for the FOMC held near the second month of the quarter.

contemporaneous unemployment gap). In the first stage regressions, the R-squared statistics are fairly large, ranging from .45 to .97, suggesting the endogenous variables are highly correlated with the instruments.

[19] The empirical work used the preliminary estimates of core PCE inflation, usually released by the end of the first month of a quarter. The Greenbook forecasts used as instruments were the ones prepared for the FOMC meetings held near the second month of a quarter. This timing means that the Board staff preparing the Greenbook forecasts had information about the preliminary estimates of core inflation rates in previous quarters. However, none of the conclusions reported here would change if we had used third release estimates, usually reported by the end of the third month of the quarter.

[20] In January of each year from 1991–2006, the CBO released estimates of the NAIRU. For the period 1987–1990, the estimates used are those given in the 1991 vintage data file. For 1991, we used the pertinent series on the NAIRU from the 1992 vintage data file and so on for each year thereafter.

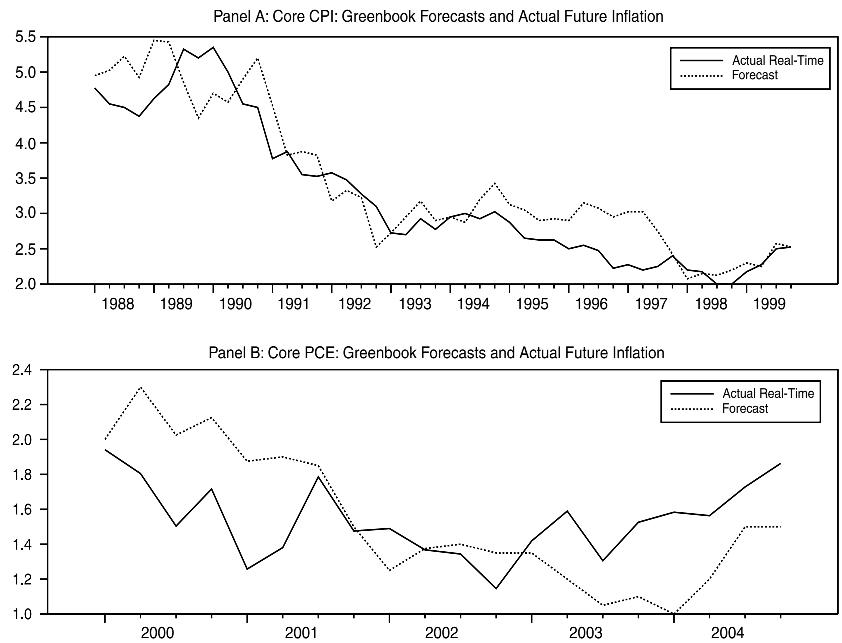
Figure 3 Unemployment Gap



Panel A in Figure 1 charts the four-quarter averages of real-time headline and core CPI inflation rates from 1988:1–1999:4, and Panel C charts the averages of headline CPI and core PCE inflation rates from 2000:1–2006:4. As can be seen, headline and core CPI inflation series stay together for most of the period before 2000 (see Panel B). However, over 2000:1–2006:4, headline CPI inflation remained above core PCE inflation (see Panel D), suggesting that a policy rule that relates the policy rate to headline CPI inflation is likely to prescribe a higher federal funds rate target than a policy rule that relates the policy rate to core PCE inflation, ceteris paribus. Hence, given the different behavior of headline CPI inflation and core PCE inflation rates over this subperiod, the measure of inflation used in the estimated Taylor rule will matter for predicting the stance of monetary policy.

Figures 2 and 3 chart real-time and 2009 vintage estimates of economic fundamentals that enter the Taylor rules; Figure 2 charts the four-quarter average of core CPI and core PCE inflation rates, whereas Figure 3 charts the unemployment gap. Two observations are noteworthy. First, core PCE inflation data have been extensively revised over the years, and there are big discrepancies between real-time and revised estimates of core PCE inflation. In particular, real-time estimates of the four-quarter average of PCE inflation

Figure 4 Greenbook Inflation Forecasts and Actual Future Inflation (4Q Average)



rates were substantially below the 2009 vintage estimates over 2002:1–2005:4 (see Figure 2, Panel B). Second, the unemployment gap data is also revised, but discrepancies between the real-time and 2009 vintage estimates are small and do not increase appreciably over 2001–2006 (see Figure 3).

Figure 4 charts Greenbook inflation forecasts and actual future inflation; Panel A charts inflation forecasts of core CPI inflation and Panel B charts those of core PCE. As can be seen, Greenbook inflation forecasts of core CPI inflation do track actual core CPI inflation, with the exception of the short period 1994–1997 when the Greenbook forecasts turned out to be too pessimistic. For the later period, the Greenbook forecasts of core PCE inflation do not fare as well in predicting actual core inflation. In particular, in 2003:1–2004:4, the Greenbook forecasts of core PCE inflation indicated deceleration in expected inflation, but actual core PCE inflation turned out to be much higher than what the Board staff predicted. The fear of expected deflation implicit in Greenbook forecasts of declining future inflation is used by some analysts to argue that the Greenspan Fed may have kept the federal funds rate target too low for too long during this subperiod. However, it is for a very short span that actual core inflation was higher than what the Board staff forecasted. The result

Table 1 Estimated Taylor Rules

Row	End Period	Inflation	α_π	α_y	ρ	s	SER
1	1999:4	Core CPI + PCE	1.6	−1.3	.56	.62	.322
			(5.3)	(4.3)	(3.3)	(2.9)	
2	2004:4	Core CPI + PCE	1.9	−1.4	.52	.67	.331
			(8.2)	(5.5)	(6.4)	(3.6)	
3	1999:4	Headline CPI	1.5	−1.0	.72	.61	.352
			(2.9)	(2.3)	(5.8)	(3.2)	
4	2004:4	Headline CPI	.1	−2.4	.40	.97	.347
			(.3)	(−4.9)	(2.6)	(3.3)	
5	1999:4	Core CPI	1.6	−1.3	.56	.62	.322
			(5.4)	(4.3)	(3.3)	(2.9)	
6	2004:4	Core CPI	1.6	−1.4	.60	.75	.324
			(3.1)	(3.0)	(4.3)	(4.4)	

Notes: Rows labeled 1 through 4 contain nonlinear instrumental variables estimates of policy coefficients from the forward-looking policy rule given below in (a) and use real-time data on inflation and the unemployment gap:

$$\begin{aligned} FR_t &= \alpha_0 (1-s)(1-\rho) + (s+\rho) FR_{t-1} + (1-\rho)\left\{\alpha_\pi E_t \bar{\pi}^c_4 + \alpha_u \left(ur_t - ur^*_t\right)\right\} \\ &\quad -s\left\{(1-\rho)\alpha_\pi E_{t-1}\bar{\pi}^c_4 + (1-\rho)\alpha_u\left(ur_{t-1} - ur^*_{t-1}\right)\right\} - s\rho FR_{t-2} + \varepsilon_t. \quad \text{(a)} \end{aligned}$$

The instruments used are three lagged values of Greenbook inflation forecasts, the funds rate, unemployment gap, the growth gap, and the spread between nominal yields on 10-year Treasury bonds and the federal funds rate. Parentheses contain t-values. SER is the standard error of estimate. Estimation was done allowing for the presence of first-order serial correlation in v_t, and s is the estimated first-order serial correlation coefficient. The sample periods begin in 1988:1 and end in the year shown in the column labeled "End Period."

here—that a rational expectations version of the Taylor rule estimated using real-time data tracks the actual funds rate target well—implies that the fear of deflation may have played a limited role in keeping the funds rate target low during this subperiod.

2. EMPIRICAL RESULTS

This section presents and discusses policy response coefficients from Taylor rules that are estimated using different measures of inflation. It also examines the stability of policy response coefficients using the Chow test with the break data around 2000, when the Greenspan Fed switched from focusing on CPI to PCE inflation measure.

Estimates of Policy Response Coefficients

Table 1 presents estimates of policy response coefficients (α_π, α_u) from the Taylor rule in equation (4) for two sample periods, 1988:1–1999:4 and

1988:1–2004:4. Rows 1 and 2 present estimates using the time-varying measure of core inflation, and rows 3 and 4 present estimates for headline CPI inflation measure. Focusing first on estimates of the Taylor rule with the time-varying measure of core inflation, all estimated policy response coefficients are correctly signed and statistically significant. In particular, the inflation response coefficient α_π is generally well above unity, suggesting that the Greenspan Fed responded strongly to expected inflation. Furthermore, in both sample periods, estimated policy response coefficients remain correctly signed and are statistically significant, suggesting parameter stability.[21,22]

Focusing on estimates of the Taylor rule with headline CPI inflation, we find that estimated policy response coefficients are sensitive to the sample period. For the sample period ending in 1999:4, the estimated policy response coefficients are correctly signed and statistically significant. The estimated inflation response coefficient is 1.5, well above unity, and the estimated unemployment gap response coefficient is close to unity. However, the estimated policy response coefficients are not stable across the two sample periods. In particular, the estimated inflation response coefficient falls below unity and is no longer statistically significant when the policy rule is estimated over 1988:1–2004:4 (see Table 1, Row 4). This result is similar in spirit to the one in Smith and Taylor (2007), who estimate a Taylor rule over 1984:1–2005:4 and find that the estimated inflation response coefficient declined significantly in 2002, leading them to conclude that the Greenspan Fed had become less responsive to inflation.

Parameter Stability

We formally test for stability of policy response coefficients in the Taylor rule over 1988:1–2004:4 using the Chow test and treating the break date as unknown. Since the FOMC switched to the PCE measure of inflation in

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[21] Other estimated coefficients of interest are also correctly signed. The estimated serial correlation coefficient, s, is generally positive and statistically significant, indicating the presence of serially correlated errors in the estimated policy rules. As noted in Rudebusch (2006), the presence of serial correlation may reflect influences on the policy rate of economic variables to which the Federal Reserve may have responded but that are omitted from the estimated policy rule. Furthermore, even after allowing for the presence of serial correlation, the estimated partial adjustment coefficient, ρ, is positive and well above zero, suggesting that the continued role of partial adjustment in generating a significant coefficient on the lagged value of the funds rate. This result is in line with the one in English, Nelson, and Sack (2002). However, the magnitude of the estimated partial adjustment coefficient, ρ, reported here is somewhat smaller than what is found in previous research.

[22] The empirical work employed the inflation series using CPI until 2000:4 and PCE thereafter. The estimates of the policy response coefficients do not change much if the policy rule is alternatively estimated using CPI until 2000:1 and PCE thereafter. Furthermore, the test of parameter stability discussed in the next section was implemented for all break dates over 2000:1–2001:4. As discussed later, the estimated policy rule employing the time-varying measure of inflation did not indicate a break in policy response coefficients for any of the break dates.

Table 2 Test for Stability of Policy Coefficients in Policy Rules

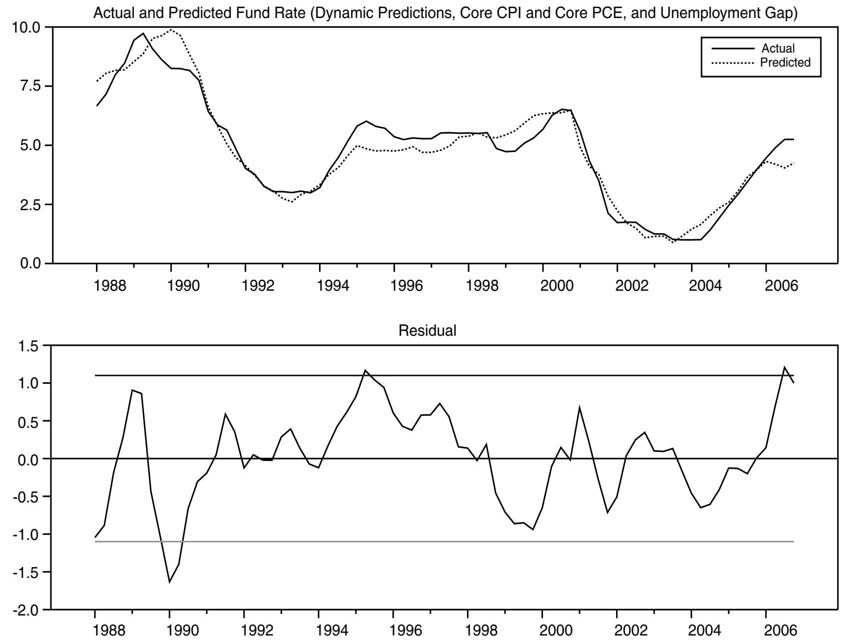
Breakpoint	**Policy Rule Core CPI + Core PCE (A)**	**Policy Rule Headline CPI (B)**	**Policy Rule Core CPI (C)**
2000Q1	.86	.04	.22
2000Q2	.95	.02	.19
2000Q3	.46	.01	.16
2000Q4	.30	.01	.17
2001Q1	.41	.00	.05
2001Q2	.17	.00	.01
2001Q3	.65	.00	.19
2001Q4	.75	.00	.35

Notes: The values reported are p-values of a test of the null hypothesis in which policy coefficients, including the intercept in the policy rule, were stable against the alternative in which coefficients changed at the indicated date. Since the test is implemented including dummy variables in the policy rule given in equation (a) in the Table 1 Notes, the reported p-values are a test of the null hypothesis in which coefficients on slope dummies, including the intercept, did not change at the indicated date.

2000, we look for a break in the estimated Taylor rule around that period. In particular, for each date between 2000:1–2001:4, we include intercept and slope dummies on policy response coefficients in the Taylor rule in equation (4) and test their joint significance for a possible break in the estimated relation. Table 2 reports the p-value for a test of the null hypothesis in which Taylor rule coefficients were stable against the alternative in which coefficients changed at the indicated date. The column labeled (A) reports p-values generated using the Taylor rule that employed the time-varying measure of core inflation, whereas the column labeled (B) does so for the Taylor rule with headline CPI inflation. As can be seen, there is no date in the interval 2000:1–2001:4 at which one could claim to find a statistically significant break in the Taylor rule if one uses a time-varying measure of core inflation. In contrast, there are several dates one could find the evidence of a break in relation if the Taylor rule is estimated using headline CPI inflation (see column B). The latter result is similar in spirit to the one in Smith and Taylor (2007).[23]

[23] The test for parameter stability was implemented using intercept and slope dummies. In the case of the policy rule that was estimated using the time-varying measure of inflation, both the intercept and slope dummy coefficients were not different from zero, suggesting that there was no shift in the intercept of the policy rule in response to change in the measure of inflation employed. In contrast, when the policy rule is estimated using headline CPI, the slope dummy coefficient on the inflation response coefficient is relatively small, suggesting that the Federal Reserve did not respond as aggressively to headline inflation as it did before. This result is in line with the inflation response coefficient becoming insignificant when the policy rule is estimated over 1988:1–2004:4 (compare estimates across rows 3 and 4, Table 1).

Figure 5 Forward-Looking Taylor Rule



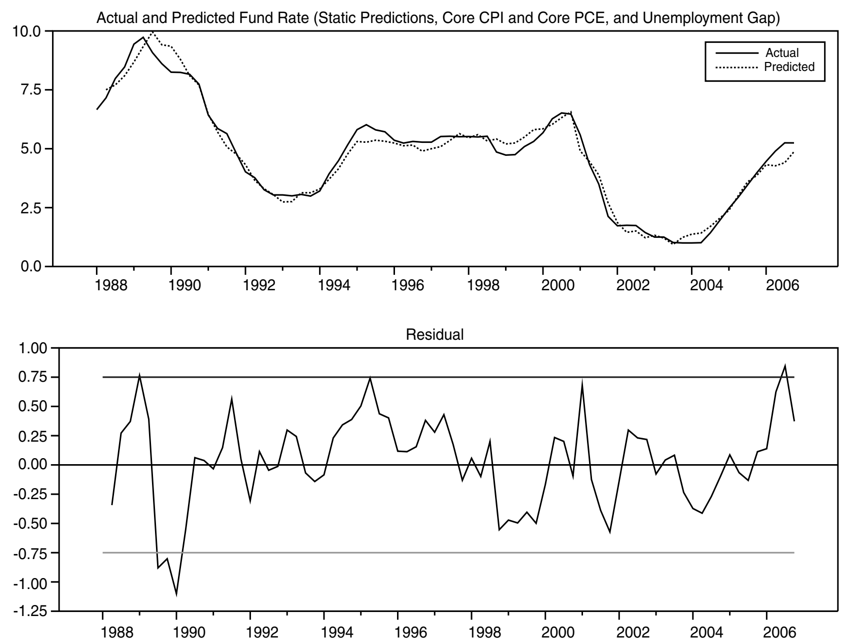
Predicting the Policy Rate Using a Taylor Rule based on Core Inflation: Was the Fed Off the Taylor Rule over 2001:1–2006:4?

In order to evaluate whether monetary policy actions over 2000:1–2006:4 can be explained by a Taylor rule, we generate predictions of the policy rate using the estimated Taylor rules. We consider two Taylor rules that differ with respect to the measure of inflation, and we generate both dynamic and static predictions. The dynamic predictions are generated using the policy rule as shown in (5):

$$FR_t^p = \hat{\rho} FR_{t-1}^p + \left(1 - \hat{\rho}\right) \left\{\hat{\alpha}_0 + \hat{\alpha}_\pi \bar{\pi}^c_{t,\bar{4}} + \hat{\alpha}_u \left(ur_t - ur_t^*\right)\right\}, \tag{5}$$

where FR^p is the predicted funds rate and the other variables are defined as before. As can be seen in the prediction equation given in (5), in generating the current quarter predicted value of the funds rate, we use last quarter's predicted value of the federal funds rate rather than the actual value, while using current-period values of the other two economic fundamentals. As a result, the current funds rate is a distributed lag on current and past values of expected inflation and the unemployment gap.

Figure 6 Forward-Looking Taylor Rule



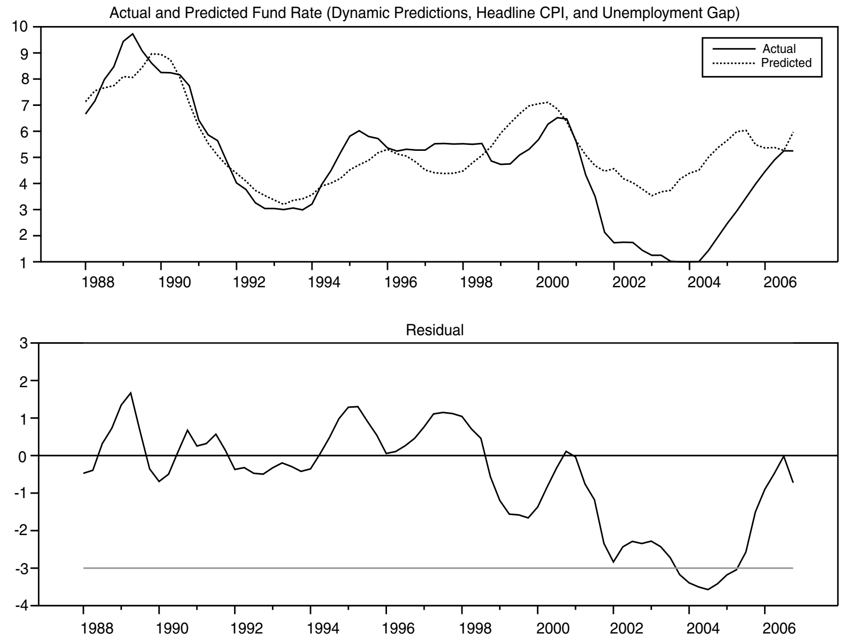
In contrast, the static predictions of the policy rate are generated while also paying attention to recent policy actions, in addition to economic fundamentals. In particular, the static predictions are generated using the estimated policy rule as shown in (6):

$$FR_t^p = \hat{\rho} FR_{t-1} + \left(1 - \hat{\rho}\right) \left\{ \hat{\alpha}_0 + \hat{\alpha}_\pi \bar{\pi}^c_{t,\bar{4}} + \hat{\alpha}_u \left(ur_t - ur_t^* \right) \right\}. \qquad (6)$$

The policy rule shown in (6) is similar to the one in (5) with the exception that (6) uses last quarter's actual value of the federal funds rate. Thus, in the static exercise the current forecast is influenced in part by actual policy actions, the magnitude of the influence of policy on the forecast being determined by the size of the partial adjustment coefficient, ρ.[24]

[24] Since the dynamic predictions are generated by paying attention only to expected inflation and the unemployment gap, they are better at revealing certain types of misspecification. In particular, if the federal funds rate equation is misspecified because it is estimated ignoring the influences of some other economic fundamentals, then the dynamic predictions generated using such a policy rule are likely to be poor proxies for the actual behavior of the federal funds rate. Hence, the dynamic predictions are better at gauging the fit of the estimated policy rule than are the static predictions.

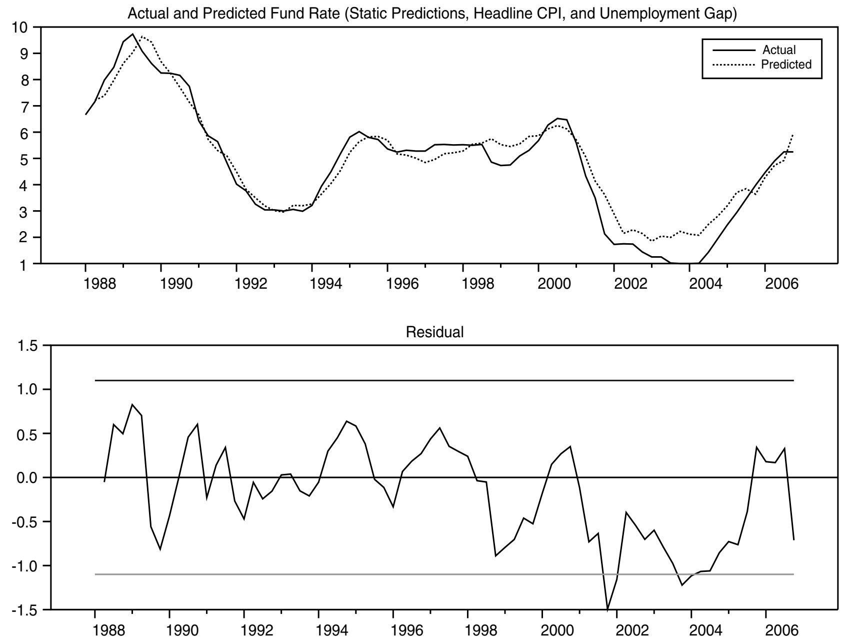
Figure 7 Forward-Looking Taylor Rule



Figures 5 and 6 respectively chart the dynamic and static predictions of the funds rate from the Taylor rule that is estimated using the time-varying measure of core inflation.[25] Actual values of the funds rate and the prediction errors are also charted there. Two observations need to be highlighted. First, the estimated policy rule predicts very well the broad contours of the policy rate over 1988:1–2006:4. The mean absolute error is .47 percentage points when dynamic predictions are used and .30 percentage points when static predictions are used. The root mean squared error is .60 percentage points when dynamic predictions are used, whereas it is only .38 percentage points when static predictions are used. Secondly, focusing on the period from 2000:1–2006:4, there is no evidence of persistently large prediction errors, and most prediction errors are small in magnitude (below twice the root mean squared error), suggesting that the actual funds rate is well predicted and, hence, that the Greenspan-Bernanke Fed was "on" a Taylor rule.

[25] The predictions begin in 1988:1. For generating the prediction for 1988:2, we use last quarter's actual funds rate. For later periods, the predicted values are generated using last period's predicted value and current period estimates of expected inflation and the unemployment gap.

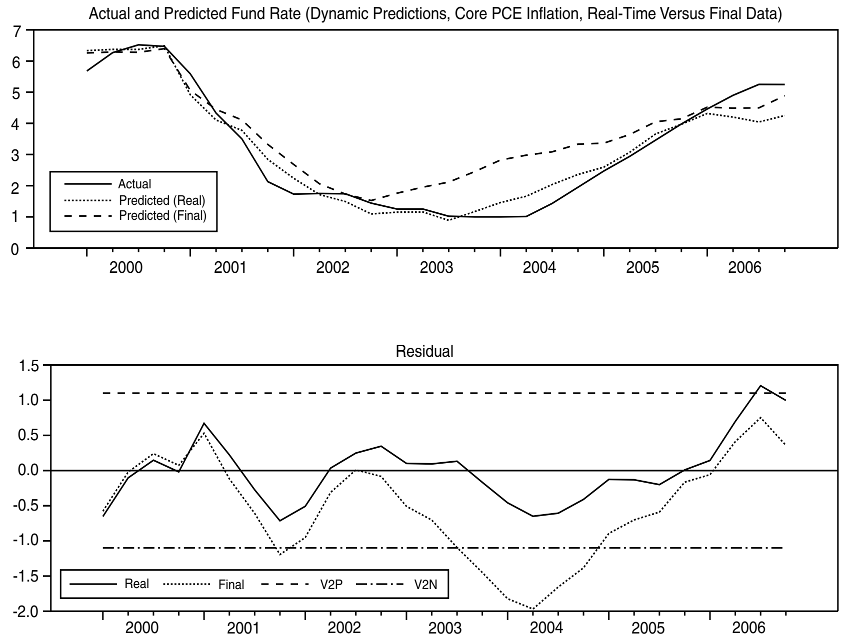
Figure 8 Forward-Looking Taylor Rule



Predicting the Policy Rate Using a Taylor Rule Based on Headline CPI Inflation: Was the Fed Off the Taylor Rule over 2001:1–2006:4?

Figures 7 and 8 respectively chart the dynamic and static predictions of the policy rate from the Taylor rule estimated using headline CPI inflation. Two observations are noteworthy. First, this Taylor rule does not predict well the broad contours of the policy rate. The mean absolute error is 1.1 percentage points and the root mean squared error is 1.5 percentage points, based on dynamic prediction of the policy rate. The summary measures of predictive performance improve somewhat when they are calculated using the static prediction errors—the mean absolute error is .46 percentage points and the root mean squared error is .56 percentage points. Secondly, focusing on the period from 2000:1–2006:4, there is clear evidence of persistently large negative prediction errors, and many of these prediction errors are large in magnitude (see lower panels, Figures 7 and 8). According to this Taylor rule, the actual funds rate remained consistently below the level prescribed, implying policy was too loose for most of the period over 2000:1–2006:4—a result that is in line with the ones in Smith and Taylor (2007) and Taylor (2007).

Figure 9 Counterfactual Simulations



Role of Data Revisions

Figure 2 shows that data on core PCE inflation have been extensively revised over the years, particularly for the period 2002–2005 when real-time estimates of core PCE inflation are substantially below the 2009 vintage estimates. Figures 9 and 10 chart, respectively, the counterfactual dynamic and static simulations of the policy rate generated using 2009 vintage data on economic fundamentals.[26] For a comparison, the predictions generated using real-time data are also charted. As can be seen, deviations of the policy rule using the 2009 vintage data are somewhat larger than those generated using real-time data. However, it would be misleading to conclude from such evidence that the Federal Reserve was too loose.[27]

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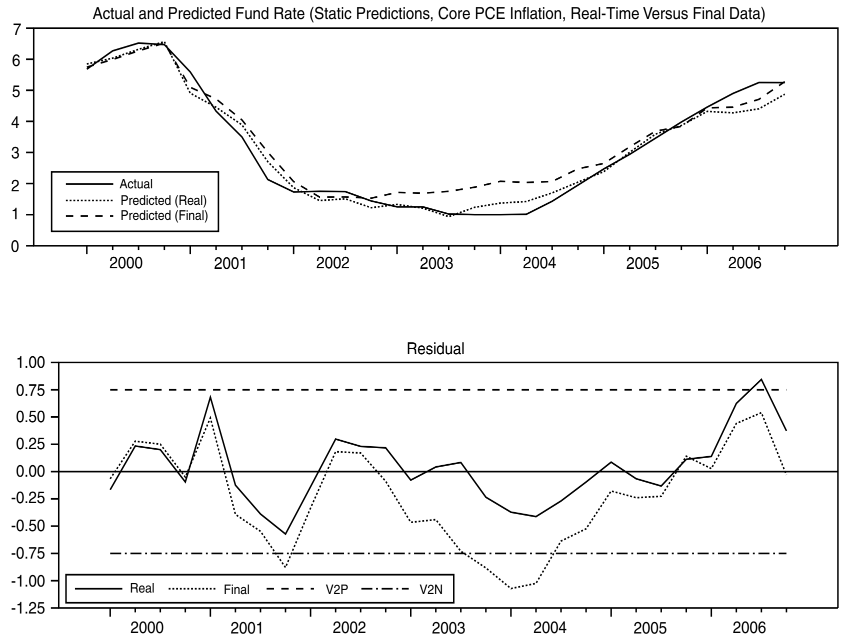
[26] The policy rule is estimated using real-time data over 1988:1–2004:4. The dynamic predictions are, however, generated using not real-time but 2009 vintage estimates of core PCE inflation and the unemployment gap.

[27] Many analysts have examined other indicators of inflation available in real time and conclude that monetary policy was not inflationary, despite the low level of the federal funds rate target. For example, Dokko et al. (2009) have examined the commercially available inflation forecasts of the private sector as well as the inflation forecasts made by the individual members the FOMC published in the Humphrey-Hawkins reports over 2003–2006. They concluded that all those inflation forecasts were consistent with the Federal Reserve's informal inflation target of between

Figure 10 Counterfactual Simulations

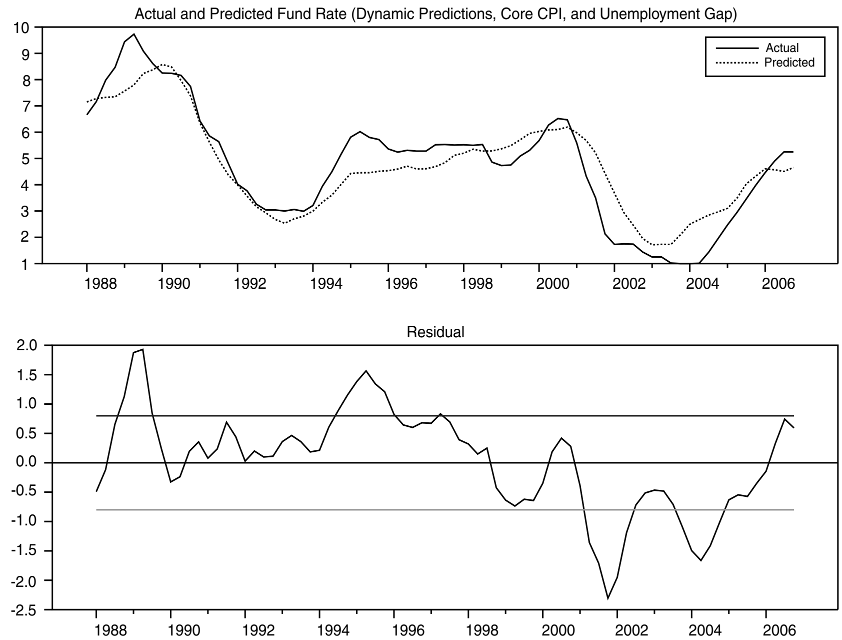


Role of Deflation Fears

Some analysts, focusing on the Taylor rule estimated using CPI inflation measure, contend that, over the period 2002:1–2005:4, the Greenspan Fed may have kept the federal funds rate too low for too long in order to avoid the consequences of a Japanese-style deflation. According to this explanation, internal forecasts of the U.S. inflation rate indicated the possibility of deflation, which led the Greenspan Fed to keep the short-term interest rate low for an extended period of time. There is some limited support for this view in Figure 4, which shows that the Greenbook forecasts of core PCE inflation indicated substantial deceleration of expected inflation for most of the period over 2002:1–2005:4. However, actual core PCE inflation did not decline to levels indicated by the Greenbook forecast. Also, as shown above, the actual funds rate is close to what is prescribed by a forward-looking Taylor rule estimated using real-time data on the fundamentals, namely, core PCE inflation

1.5 percent to 2 percent. Others focusing on the bond market measures of inflationary expectations point out that, over this subperiod, long-term rates exhibited considerable stability that is consistent with the presence of a noninflationary policy stance, despite the low level of the federal funds rate.

Figure 11 Forward-Looking Taylor Rule

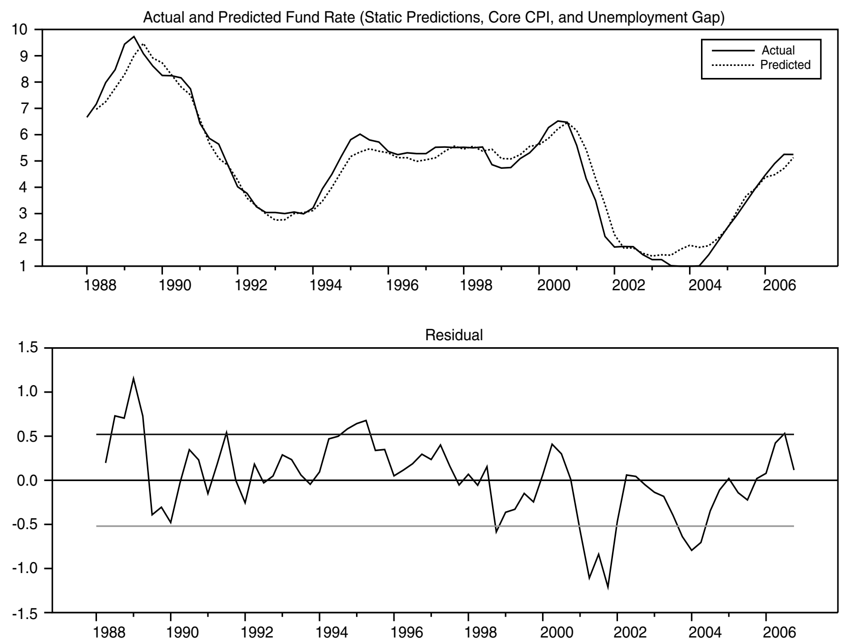


and the unemployment gap. The empirical work here suggests that, while the fear of deflation may have played some role, the actual funds rate remained low for fundamental reasons once we recognize that the Greenspan Fed was focused not on headline CPI but on a core measure of PCE inflation.

Headline CPI Versus Core CPI

The result here—that a forward-looking Taylor rule estimated using a headline measure of CPI inflation does not depict parameter stability during the Greenspan years—continues to hold if the Taylor rule is instead estimated using a core measure of CPI inflation. In fact, several analysts, including Blinder and Reis (2005), have estimated Taylor rules using a core measure of CPI inflation. But, as shown below, the use of a core measure of CPI inflation does generate reasonable estimates of policy response coefficients; the estimated policy rule, however, does not depict parameter stability in the Greenspan years. Table 1 presents policy response coefficients estimated using core CPI inflation data and Table 2 presents p-values of the Chow test of parameter stability (see column C). As can be seen, estimated policy response coefficients appear reasonable. However, the estimated policy rule still

Figure 12 Forward-Looking Taylor Rule

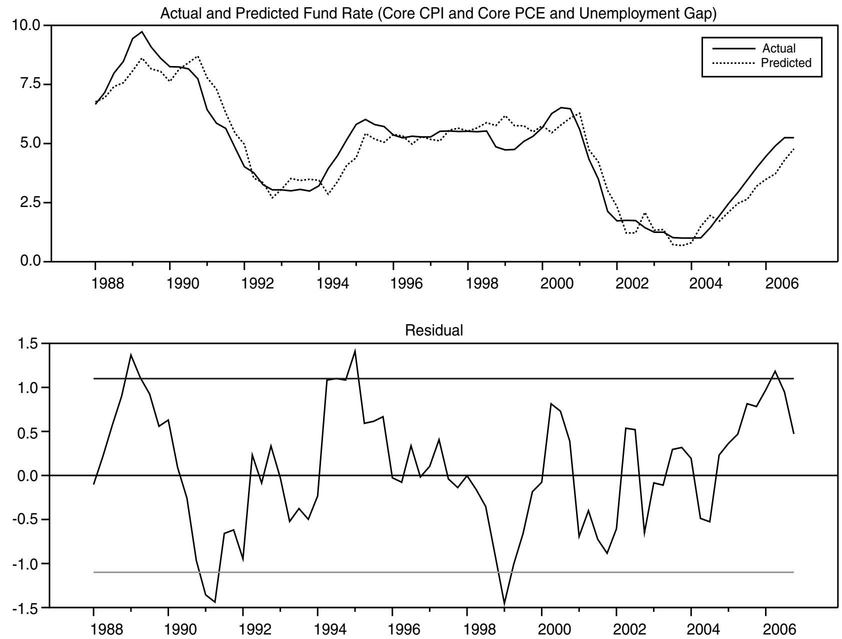


exhibits parameter instability in 2001; the test results indicate a reduction in the size of the inflation response coefficient, consistent with the observation in Taylor (2007) that the Greenspan Fed did not react strongly to inflation after 2001. Figures 11 and 12 chart the dynamic and static simulations of the federal funds rate using the estimated Taylor rule based on the core measure of CPI inflation. As can be seen, the actual funds rate is considerably below the value prescribed by this policy rule for most of the subperiod from 2001:1–2006:4. Using the metric of summary error statistics based on dynamic predictions, we calculate the mean absolute error as .70 percentage points and the root mean squared error as .86 percentage points. By this metric, the Taylor rule estimated using core CPI does better than the Taylor rule estimated using headline CPI inflation. However, neither of these Taylor rules depict parameter stability and both are consistent with policy by being "too loose" over most of the period 2002:1–2006:4.

Forward- Versus Backward-Looking Taylor Rules

The empirical work here has used forward-looking Taylor rules to show that the measure of inflation chosen matters for predicting actual policy actions

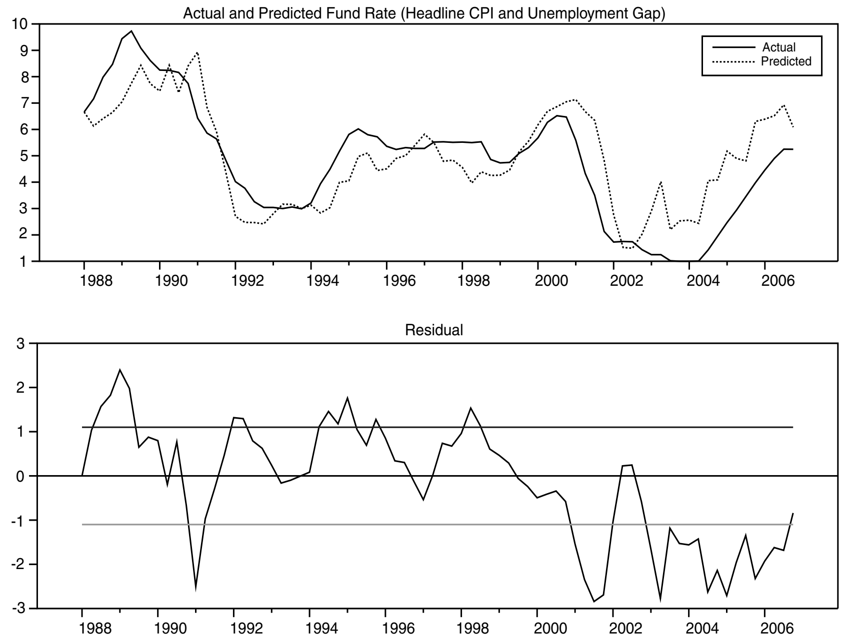
Figure 13 Backward-Looking Taylor Rule



over 2002:1–2006:4, namely, a Taylor rule estimated using the time-varying measure of core inflation tracks actual policy better than a Taylor rule estimated using headline CPI inflation. However, it may be noted that the above result continues to hold if one estimates and compares backward-looking Taylor rules. Namely, a backward-looking Taylor rule estimated using the time-varying measure of core inflation tracks actual policy actions much better than does a Taylor rule with headline CPI inflation (see Figures 13 and 14). However, backward-looking Taylor rules generally do not depict parameter stability, even when they are estimated using the time-varying measure of core inflation.[28]

[28] The Taylor rules considered in this exercise were estimated using smoothed lagged values of inflation and unemployment gap variables, as in the original Taylor rule. We also estimated versions in which we include a lagged value of the federal funds rate, thereby directly allowing interest rate smoothing. This specification gave qualitatively similar results.

Figure 14 Backward-Looking Taylor Rule



Discussion of Results

The empirical results above suggest that monetary policy actions in the Greenspan era can be summarized by a stable Taylor rule according to which the Federal Reserve was forward looking, smoothed interest rates, and focused on a core measure of inflation measured by CPI until 2000 and PCE thereafter. The estimated Taylor rule does not depict any parameter instability, despite the switch in the measure of inflation used in monetary policy deliberations. In contrast, Taylor rules that do not allow for this switch in the measure of inflation, and are instead estimated using CPI inflation (headline or core), depict parameter instability around 2000, indicating that the Greenspan Fed did not react strongly to expected (CPI) inflation.

Within the context of a Taylor-type policy rule, a switch in the measure of inflation is likely to affect the policy rule mainly by altering the constant term of the policy rule if the switch leads to a different inflation target expressed in a new inflation measure. This occurs because the constant term in a Taylor rule has embedded in it the Fed's estimate of its inflation target. However, the constant term in a Taylor rule also has embedded in it the Fed's estimate of the economy's real rate of interest. To see it, rewrite equation (1.1) as

$FR_t^* = rr^* + \pi^* + \alpha_\pi E_t\left(\pi_{t+j}^c - \pi^*\right) + \alpha_u\left(ur_t - ur_t^*\right)$, where rr^* is the real rate and π^* is the inflation target. If we substitute the above equation into equation (1.2), we get equation (1.3), where the constant term is now defined as $\alpha_0 = rr^* + (1 - \alpha_\pi)\,\pi^*$. The constant term thus has embedded in it the Fed's estimates of the real rate of interest as well as its inflation target. However, as is well known, given a reduced-form estimate of the constant term, we can't recover the Fed's estimates of the real rate as well as its inflation target without bringing in additional information.

The switch in the measure of inflation from core CPI to core PCE does not appear to be associated with any significant shift in the estimated Taylor rule used to explain monetary policy actions in the Greenspan years.[29] One possible explanation of why the switch did not lead to any significant shift in the estimated Taylor rule is that while the switch may have lowered the Fed's inflation target expressed in core PCE inflation, it may have also caused the Greenspan Fed to raise its assessment of the economy's real rate of interest, thereby leaving the constant term of the estimated Taylor rule unchanged.[30]

3. CONCLUDING OBSERVATIONS

This article shows that the measure of inflation used in estimating Taylor rules to explain historical monetary policy actions is not innocuous. The FOMC under the chairmanship of Alan Greenspan refined the measure of inflation used in monetary policy deliberations, switching from focusing on CPI to focusing on PCE in the early 1980s. Moreover, Chairman Greenspan encouraged both the FOMC and the financial markets to focus on core rather than headline inflation in implementing policy. As noted in Blinder and Reis (2005), during the Greenspan era an oil shock was considered a "blip" in the inflation process that did not affect long-term inflationary expectations and therefore should be ignored, leading the Fed to focus on core rather than headline inflation in the implementation of monetary policy.

If we estimate a Taylor rule that uses real-time data and we employ the time-varying measure of core inflation, then the estimated policy rule depicts

[29] This result is consistent with the test results of parameter stability discussed above. For each possible break date between 2000:1–2001:4, the Chow test of parameter stability was performed including intercept as well as slope dummies on response coefficients in the policy rule. For all the break dates, the intercept dummy was not statistically different from zero, which is consistent with the absence of a change in the constant term of the policy rule.

[30] Using the metric of comparing means, the sample mean of core PCE inflation rates over 1987:1–2005:4 is 2.5 percent, which is lower than the value (3.1 percent) computed using core CPI inflation rates over the same period. Given the differential trend behavior of these two inflation measures, the Greenspan Fed having an inflation target of, say, 2 percent based on the behavior of core PCE inflation is equivalent to its having an inflation target of 2.6 percent based on the core CPI inflation measure. Hence, the switch from CPI to PCE measure of inflation could have been associated with a downward shift in the constant term of the estimated Taylor rule around 2000.

parameter stability in the Greenspan era and predicts very well the actual path of the federal funds rate over 2001:1–2006:4. In contrast, a Taylor rule that is estimated using headline CPI inflation depicts parameter instability and indicates the actual funds rate was too low relative to the level prescribed, as headline CPI inflation remained above core PCE inflation during most of this short period. Hence, in evaluating monetary policy actions in the Greenspan era, it is important to pay attention to these two real-time features of monetary policy deliberations, namely, the focus on core rather than headline inflation measures and the switch from CPI inflation to PCE inflation.

Following John Taylor's (2007) work, many analysts and some policymakers have begun to contend that, over 2002:1–2005:4, the Federal Reserve may have lowered the federal funds rate too low for too long, suggesting that monetary policy was too loose as seen through the lens of a Taylor rule. The popular explanation of this easier stance on monetary policy during this period is that the Greenspan Fed feared deflation. In fact, the Greenbook forecasts of core PCE inflation indicated substantial deceleration in expected inflation during this subperiod, which did not materialize. However, the result here—that a forward-looking Taylor rule estimated using real-time core PCE inflation data tracks the actual funds rate well—implies that the actual funds rate was determined for fundamental reasons. In real time, the Fed's preferred measure of core PCE inflation fluctuated in a low narrow range.

The core measure of PCE inflation has been extensively revised over the years. In particular, the most recent 2009 vintage data indicates that over 2002–2006, core PCE inflation did not decelerate as much and was substantially higher than what the Federal Reserve knew in real time. When seen through the lens of a Taylor rule, policy deviations using the 2009 vintage data are somewhat larger than those generated using the real-time data. However, it would be misleading to conclude from such evidence that monetary policy was too easy. Several other indicators of inflationary expectations that were available in real time indicate policy was noninflationary over this subperiod.

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Monetary Policy with Interest on Reserves

Andreas Hornstein

In response to the emerging financial crisis of 2008, the Federal Reserve decided to increase the liquidity of the banking system. For this purpose, the Federal Reserve introduced or expanded a number of programs that made it easier for banks to borrow from it. For example, commercial banks were able to obtain additional loans through the Term Auction Facility, which the banks would then hold in their reserve accounts with the Federal Reserve. As a result of the combined financial market interventions, the balance sheet of the Federal Reserve increased from about \$800 billion in September 2008 to more than \$2 trillion in December 2008. Over the same time period, the reserve accounts of commercial banks with the Federal Reserve increased from about \$100 billion to \$800 billion. In late 2008 the Federal Reserve also announced a program to purchase mortgage-backed securities (MBS) and debt issued by government-sponsored agencies. Since then, outright purchases of agency MBS and agency debt have essentially replaced short-term borrowing by commercial banks on the asset side of the Federal Reserve's balance sheet, and the volume of outstanding reserves increased again to about \$1.1 trillion by the end of 2009. Given the magnitude of outstanding reserves, there is some concern these reserves might limit policy options once the Federal Reserve decides to pursue a more restrictive monetary policy. Yet, another change in the available policy instruments might lessen this concern: Starting in October 2008, the Federal Reserve began to pay interest on the reserve accounts that banks hold with the Federal Reserve System.

How should one think about monetary policy when reserve accounts earn interest? To study this issue, I introduce a stylized banking sector into a simple baseline model of money that is at the core of much research in monetary

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economics. In this framework I address an admittedly rather narrow theoretical question, but this question is fundamental to any theory of monetary policy. Namely, does the payment of interest on reserves affect issues of price level determinacy? An indeterminate price level might be undesirable since it can give rise to price level fluctuations driven by self-fulfilling expectations. In this context it is shown that the amount of outstanding reserves has only limited implications for the conduct of monetary policy.

Price level determinacy is studied in a theoretical framework that specifies not only monetary policy, but also fiscal policy, e.g., Leeper (1991) or Sims (1994). Monetary policy is described as setting a short-term nominal interest rate in response to inflation, and fiscal policy is described as setting the primary surplus in response to outstanding government debt. For the baseline monetary model without a banking sector, one obtains price level determinacy if monetary policy is active, that is, it responds strongly to the inflation rate, and fiscal policy is passive, that is, it responds strongly to government debt.[1] Price level determinacy is also obtained when monetary policy is passive and fiscal policy is active. For the modified model with a banking sector, I find that this characterization of price level determinacy is not materially affected, whether or not interest is paid on reserves. I obtain a determinate price level when monetary policy is sufficiently active and fiscal policy is sufficiently passive, or vice versa. Furthermore, the magnitude of outstanding reserves may not matter at all, and if it does matter the impact of reserves is small.

Earlier theoretical work on paying interest on reserves was concerned that this policy would lead to price level indeterminacy. Sargent and Wallace (1985) argue that, depending on how interest on reserves is financed, an equilibrium might not exist or the price level might be indeterminate.[2] In terms of the above characterization of monetary and fiscal policy, these results obtain because the assumed financing schemes for interest on reserves make monetary and fiscal policy both passive or both active. My results are in line with the recent work of Sims (2009), who studies the monetary and fiscal policy coordination problem when interest is paid on money in a baseline monetary model. The results are also related to Woodford's discussion (2000) of monetary policy as an interest rate policy in environments where the role of money is diminished over time.

The framework of this article is not suited to address the question of whether interest on reserves allows a separation of monetary policy from credit policy as suggested by Goodfriend (2002) and Keister, Martin, and McAndrews (2008). In this article I use a reduced form representation of liquidity preferences by households to model distinct household demand

[1] The terminology follows Leeper (1991).

[2] Smith (1991) raises similar concerns on price level determinacy in an extended version of the environment studied by Sargent and Wallace (1985).

functions for cash, bank demand deposits, and government bonds, but the model of the financial system's attitude toward the liquidity of assets in the financial system is even more rudimentary. First, I do not allow for credit risk; and second, the banks' attitudes toward liquidity risk are captured by one exogenous parameter, the desired ratio of liquid assets to deposits. The fact that the volume of reserves is of only limited relevance for price level determinacy therefore does not say anything about the ability of reserves to enhance the liquidity of the financial system.

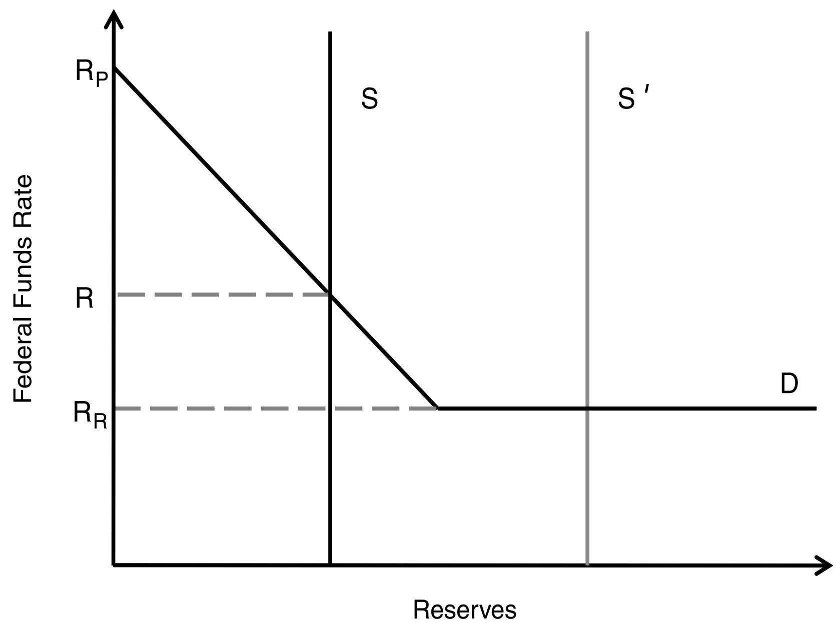
The analysis of the conduct of monetary policy when interest is paid on reserves is based on a stylized model of the economy. Before proceeding with this analysis I will review the mechanics of the Federal Reserve's interest rate policy in the next section. This section also provides an opportunity to describe how the interventions of the Federal Reserve in financial markets in 2008 affected its ability to conduct interest rate policy. Section 2 then reviews Leeper's joint analysis (1991) of monetary and fiscal policy in a baseline monetary model, and Section 3 adds a stylized banking sector to the baseline monetary model. The banking model of this section introduces the payment of interest on reserves into a simplified version of the environment studied by Canzoneri et al. (2008). Section 4 concludes.

1. THE MECHANICS OF INTEREST RATE POLICY

Most central banks implement monetary policy through an interest rate policy. That is, they target a short-term interest rate and adjust their target in response to changes in economic conditions. Federal Reserve monetary policy appears to be well-approximated by a policy rule that increases the targeted interest rate more than one-for-one in response to an increase of the inflation rate and decreases the targeted interest rate in response to declines in economic activity as measured by a declining gross domestic product growth rate or an increasing unemployment rate. This kind of behavior has become known as the Taylor rule. The short-term interest rate that the Federal Reserve targets is the federal funds rate—that is, the interest rate that U.S. banks charge each other for overnight loans. This section provides a short review of the mechanics of how the Federal Reserve influences the federal funds rate, and how paying interest on reserves affects its ability to target this rate. The review takes a very stylized view of the federal funds market, as in Ennis and Weinberg (2007). For a more detailed description of the process see Ennis and Keister (2008).

Commercial banks are required to hold particular assets (reserves) against their outstanding liabilities. How many reserves a bank has to hold depends on the types and amounts of its outstanding liabilities. Assets that qualify as reserves are vault cash and accounts with the central bank. Banks hold accounts with the central bank not only to satisfy reserve requirements, but also to facilitate intraday transactions. Private agents engage in transactions and

Figure 1 The Market for Reserves



use their bank accounts to settle payments associated with these transactions. Since not everybody is using the same bank, these payment settlements result in corresponding payment settlements between banks during a business day. Banks use their accounts with the central bank to implement these settlements. A payments transfer from one bank to another can be settled through a debit (credit) to the paying (receiving) bank's account with the central bank. Total inflows and outflows to a bank's account with the central bank during a day need not balance, and at the end of the day a bank's account may have increased or decreased. Furthermore, there is some randomness to settlement transactions and the bank is uncertain as to its end-of-day balance with the central bank.

The uncertainty about payment flows creates a problem for banks since they have to hold a certain balance with the central bank at the end of the day in order to satisfy their reserve requirement. Suppose that at the beginning of the day a bank has some amount of money and has to decide how much to allocate to its reserve account and how much to borrow/lend overnight with other banks at the federal funds rate. If the bank does not allocate enough to its reserve account and at the end of the day its balance falls short of its reserve requirement, it can borrow from the central bank at a penalty rate,

R_P.[3] Alternatively, if at the end of the day the bank's reserve account exceeds its reserve requirement, then the bank foregoes some interest income if the interest rate paid on reserve accounts, R_R, is lower than the federal funds rate.

The optimal response of banks to this settlement uncertainty creates a precautionary demand for reserves, D, that depends on the federal funds rate (Figure 1). The federal funds rate cannot exceed the penalty rate since banks can always borrow at the penalty rate. If the federal funds rate is below the penalty rate but above the interest rate paid on reserves, then the foregone interest income represents an opportunity cost to holding reserve balances. This opportunity cost, however, is declining in the federal funds rate and banks are willing to hold more reserves at lower federal funds rates. Finally, if the federal funds rate is equal to the interest on reserves, then there is no opportunity cost to holding reserves and the demand for reserves becomes infinitely elastic. The equilibrium federal funds rate is bounded by the penalty rate and the interest rate on reserves, and, given the demand for reserves, it is determined by the supply of reserves, S.

In the short run the Federal Reserve controls the federal funds rate through actions that affect the supply of reserves. The particular operating procedure for the Federal Reserve has been that the market desk at the New York Federal Reserve Bank forecasts the daily demand for reserves and then injects or withdraws reserves in order to equalize the predicted federal funds rate with the federal funds rate target set by the FOMC. Except for unusual events, the "effective" federal funds rate during the day—that is, the rate at which intrabank loans occur—is usually very close to the federal funds target rate (Figure 2a).[4] At times, when the Federal Reserve injects large amounts of liquidity for reasons other than interest rate policy, this is no longer true. For example, in response to the events of September 11, 2001, the Federal Reserve wanted to ensure the stability of the financial system and injected large amounts of reserves. This action resulted in an effective federal funds rate that was substantially below the target rate (Figure 2b). At the time, this divergence between perceived liquidity needs and interest rate policy was not considered to be a problem since the liquidity provision was viewed as temporary and to be reversed in a short period of time.

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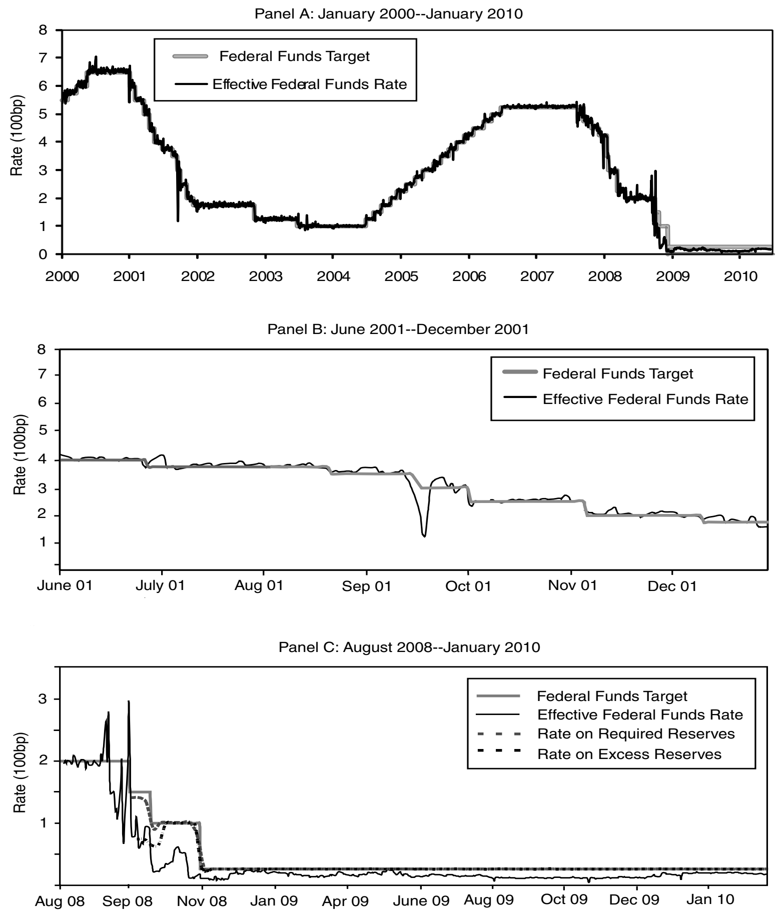
[3] In the United States, banks can borrow from the Federal Reserve against pre-approved collateral at the discount window. The discount rate is set higher than the federal funds target rate, usually 100 basis points (bp). As part of the response to the financial crisis, the Federal Reserve kept the discount rate at 25bp above the target federal funds rate from April 2008 until February 2010. A bank's effective borrowing rate is presumably higher than the discount rate since a bank's borrowing from the discount window is seen as a negative signal on the bank's balance sheet condition.

[4] Interbank lending proceeds through bilateral arrangements and, during the day, the negotiated lending rates can fluctuate substantially. The effective federal funds rate is a value-weighted average of the different loan rates.

Figure 2 Target and Effective Federal Funds Rate



Notes: The data are described in the Appendix.

After the September 2008 bankruptcy of Lehman Brothers, the Federal Reserve increased liquidity substantially in response to the widening financial crises. This was accomplished through the expansion of existing programs, such as the Term Auction Facility and swap lines to foreign central banks, and the creation of new facilities, such as the Commercial Paper Funding Facility. As a result, the Federal Reserve's balance sheet more than doubled over three months and the supply of reserves increased almost tenfold. Even if banks' demand for liquid assets increased at the time, the increase in the supply of reserves was large enough to drive the effective federal funds rate significantly below the stated federal funds target (Figure 2c).

Unlike the events of September 11, 2001, the divergence in this case between effective and target federal funds rates created a problem for the conduct of interest rate policy since the increased liquidity provision was not viewed as a short-lived measure. To deal with this problem, the Federal Reserve in October 2008 started paying interest on reserves.[5] The Federal Reserve Board initially set the interest rate on reserves below the target federal funds rate, but by early November 2008, after several modifications, the interest rate on reserves was essentially the target federal funds rate.[6]

The rationale for this policy is based on the discussion above that suggests that paying interest on reserves puts a floor to the federal funds rate (Figure 1). Thus, even if the Federal Reserve increases the supply of reserves to a point where the demand for reserves becomes infinitely elastic, e.g., S' in Figure 1, the federal funds rate should not fall below the rate paid on reserves. This suggests that with interest on reserves the Federal Reserve can separate the provision of liquidity from its interest rate policy, e.g., Goodfriend (2002). Furthermore, once the Federal Reserve pays interest on reserves, it has the choice between two policy instruments: It can continue to target a market interest rate, such as the federal funds rate, above the interest paid on reserves; or it can increase the supply of reserves sufficiently and bring the federal funds rate down to the interest paid on reserves and then adjust the interest rate it pays, e.g., Lacker (2006). The first approach targets a lending rate for banks that still contains some counterparty risk, while the second approach sets the risk-free lending rate for banks.

The actual experience with interest on reserves does not completely support this argument. Since November 2008, the effective federal funds rate has been consistently below the interest rate paid on reserves. In fact, starting in December 2008, the FOMC decided to announce a target range for the federal funds rate between 0 and 25bps. This continues to be the policy as of the writing of this article. On the positive side, since February 2008, the effective federal funds rate has traded closer to the interest rate paid on

[5] In 2006 Congress authorized the Federal Reserve to pay interest on reserves starting in 2011. At the time, the main motivation for paying interest on reserves was to eliminate the "tax distortion" implied by the absence of interest payments on reserves. For example, banks would engage in activities whose sole purpose was to minimize their holdings of "reservable" accounts.

[6] On October 6, 2008, the Federal Reserve Board announced that it would pay interest on the depositary institutions' reserve account at 10bp (75bp) below the federal funds rate target for required (excess) reserves. On October 22, the Board changed the rate paid on excess reserve balances to the lowest Federal Open Market Committee (FOMC) target rate in effect during the reserve maintenance period less 35bp. Finally, on November 5, 2008, the rate on required reserves was set equal to the average target federal funds rate over the reserve maintenance period, and the rate on excess balances was set equal to the lowest FOMC target rate in effect during the reserve maintenance period.

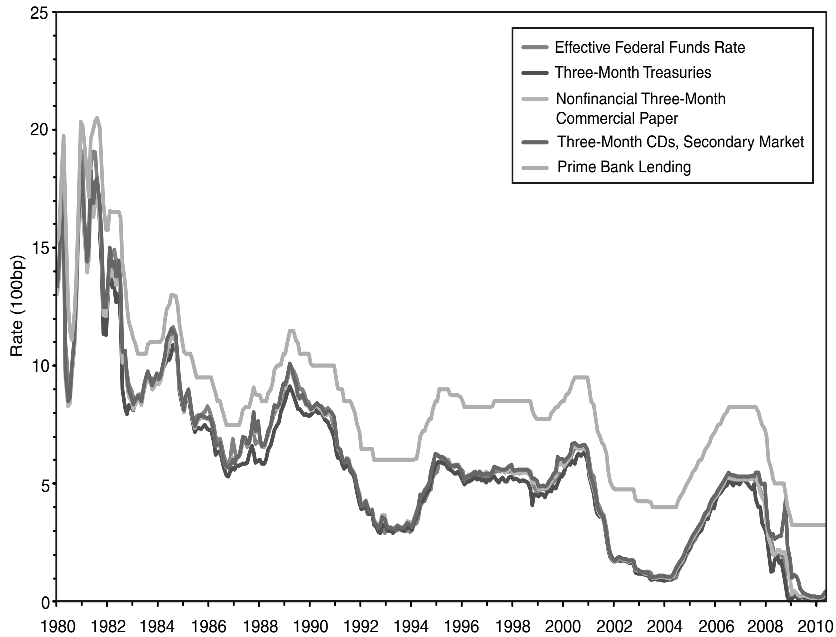
reserves. Various reasons have been advanced for this divergence between the effective federal funds rate and the interest rate on reserves. For example, in late 2008 participants in the federal funds market may have been preoccupied with events other than trying to exploit all profit opportunities in the market for overnight credit. More recently it has been argued that the low effective federal funds rate originates from particular lenders in the federal funds market—the government-sponsored enterprises (GSEs) Fannie May and Freddy Mac—who do not have interest-bearing reserve accounts with the Federal Reserve (for example, Bernanke [2009] or Bech and Klee [2010]). Arbitrage competition of depository institutions that can borrow from the GSEs and deposit the proceeds in their interest-bearing reserve accounts should eliminate any spreads between the effective rate and the reserve rate. This competition appears, however, to be limited since the GSEs apparently only engage in lending activities with a limited number of banks.

For the analysis of an interest rate policy when reserves are paying interest, I will abstract from the issues just discussed and assume that the interest rate paid on reserves is the market interest rate. First, for monetary policy I am interested in the opportunity cost to banks, which is the rate on reserves. For this analysis it is irrelevant that nonbank institutions drive the effective rate below the rate on reserves; and even if arbitrage by depositary institutions does not completely eliminate the spread between the rate on reserves and the effective rate, it will at least bound that spread. Second, for the types of aggregate models used in monetary policy analysis, there is no meaningful concept of counterparty risk. Thus, there is no risk premium that distinguishes the interbank lending rate from the riskless rate paid by reserves. Third, these models are not specified in terms of overnight interest rates, but interest rates on short-term government debt. Given that the choices for the policy rates tend to be highly persistent over short periods, this seems like a reasonable approximation. Figure 3 displays the effective federal funds rate and several other short-term interest rates from 1980 to present.[7] As is apparent from Figure 3, most of the time the different short-term interest rates track the federal funds rate quite closely.

In what follows I will study an interest rate policy that pays interest on all reserves at the market interest rate. In particular, I will study the implications of interest on reserves for price level determinacy, and to what extent the amount of outstanding reserves matters. Before proceeding to the model with interest on reserves I first outline the framework of analysis for the case without interest on reserves.

[7] All data are described in detail in the Appendix.

Figure 3 Selected Short-Term Interest Rates



Notes: The data are described in the Appendix.

2. A SIMPLE FRAMEWORK FOR THE ANALYSIS OF MONETARY AND FISCAL POLICY

The following model of an endowment economy has been used extensively in the study of monetary policy. There is one consumption good, c_t, and the consumption good is in exogenous supply. There are two nominal assets issued by the government: fiat money, M_t, and bonds, B_t. The price of the consumption good in terms of fiat money is P_t, and since the consumption good is the only good, P_t is also the price level. Inflation is the price level's rate of change from one period to the next, $\pi_{t+1} = P_{t+1}/P_t$. Bonds pay interest at the gross rate $R_{b,t}$, but fiat money does not. I define real balances and real bonds in units of the consumption good as $m_t = M_t/P_t$ and $b_t = B_t/P_t$.

Households can use both, money and bonds, to save, but holding money also provides some transactions services when households purchase consumption goods. If it was not for the transactions services, households would not want to hold money when bonds pay a positive interest rate since money does not pay any interest. The demand for real balances, equation (1), depends negatively on the opportunity cost of holding money, i.e., the foregone interest income, and positively on the real transactions volume, c_t. The demand for bonds is determined by the Euler equation (2), which equates households' willingness to exchange consumption today for consumption tomorrow with the rate at which households can do that using bonds. The latter is the real rate of return on bonds—that is, how much of the consumption good you obtain tomorrow if you invest one unit of the consumption good today in a nominal bond. Equations (1) and (2) can be derived from simple dynamic representative agent economies that explicitly specify the preferences of households and their budget constraints, e.g., Leeper (1991) or Sims (1994):

$$\begin{aligned} m_t &= \mathcal{M}\left(R_{b,t+1}\right) c_t, && (1)\\ 1 &= \beta \frac{c_t}{c_{t+1}} \frac{R_{b,t+1}}{\pi_{t+1}}, && (2)\\ v_t &= \frac{R_{b,t} v_{t-1} - \left(R_{b,t} - 1\right) m_{t-1}}{\pi_t} - \tau_t, && (3)\\ v_t &= b_t + m_t. && (4) \end{aligned}$$

Equation (3) represents the government's budget constraint. On the left-hand side is the new real debt issued to make interest payments and retire the outstanding debt on the right-hand side. Since debt is nominal, inflation reduces the real amount of debt to be repaid. Furthermore, the government does not pay interest on fiat money. Finally, if the government collects lump sum taxes, τ_t, then less new debt needs to be issued.[8] Equation (4) defines total real government debt as the sum of interest-paying real bonds and non-interest-paying real balances.

To close the model I assume that there is an exogenous endowment of the consumption good such that one can take the time path for consumption as given. I also assume that monetary policy chooses the nominal interest rate in response to the inflation rate, and fiscal policy chooses taxes in response to outstanding real bonds,

$$R_{b,t+1} = f\left(\pi_t\right) \text{ and } \tau_t = g\left(b_t\right). \qquad (5)$$

I characterize the equilibrium time paths for inflation, the interest rate, real balances, real bonds, real debt, and lump sum taxes, $x_t = \left(\pi_t, R_{b,t}, m_t, b_t, v_t, \tau_t\right)$. An equilibrium is then a bounded time path for the variables $\{x_t\}$ that solves the dynamic system defined by equations (1)–(5).[9]

[8] A negative lump sum tax represents a transfer payment to the household. We can interpret lump sum taxes as the government's primary surplus, that is, lump sum tax revenues minus spending net of interest payments.

[9] The equilibrium time paths for real balances and debt have to remain bounded, since they represent solutions to a dynamic optimization problem. Technically, real balances and debt have to satisfy transversality conditions, which state that the limiting value of the discounted future

Monetary policy is said to be active (passive) if the nominal interest rate responds strongly (weakly) to an increase of the inflation rate. Fiscal policy is said to be active (passive) if lump sum taxes respond weakly (strongly) to an increase of real bonds. For a local approximation of the difference equation system, Leeper (1991) shows that for positive interest rates there exists a unique equilibrium if monetary policy is active and fiscal policy is passive, or conversely if monetary policy is passive and fiscal policy is active.[10] The existence of a unique equilibrium in terms of the inflation rate and real balances implies price level determinacy. If both policies are passive then the equilibrium is indeterminate, and if both policies are active an equilibrium will not exist.[11] Sims (1994) shows that these results hold globally in Leeper's model (1991), and not only for local approximations.

The point of this analysis is that price level determinacy is jointly determined by monetary and fiscal policy. To illustrate this point, Figure 4, Panel A1 displays the different regions that characterize equilibrium in terms of the responsiveness of monetary and fiscal policy to the inflation rate and real debt for a standard parameterization of the model.[12] The horizontal axis displays the elasticity of lump sum taxes with respect to real debt, γ, and the vertical axis displays the elasticity of the nominal interest rate with respect to the inflation rate, α. The northeast and southwest regions represent parameter combinations for which there exist unique equilibria. The southeast region represents parameter values in which a continuum of equilibria exists, and the northwest region represents parameter values in which no equilibrium exists.

The intuition for this decomposition of the policy parameter space is fairly straightforward. Substituting the interest rate policy rule (5) into the Euler equation (2) shows that the difference equation describing the dynamics of inflation is independent of fiscal policy. If monetary policy is active, i.e., it responds strongly to past inflation, then this difference equation defines a unique bounded solution for inflation. Furthermore, if fiscal policy is passive,

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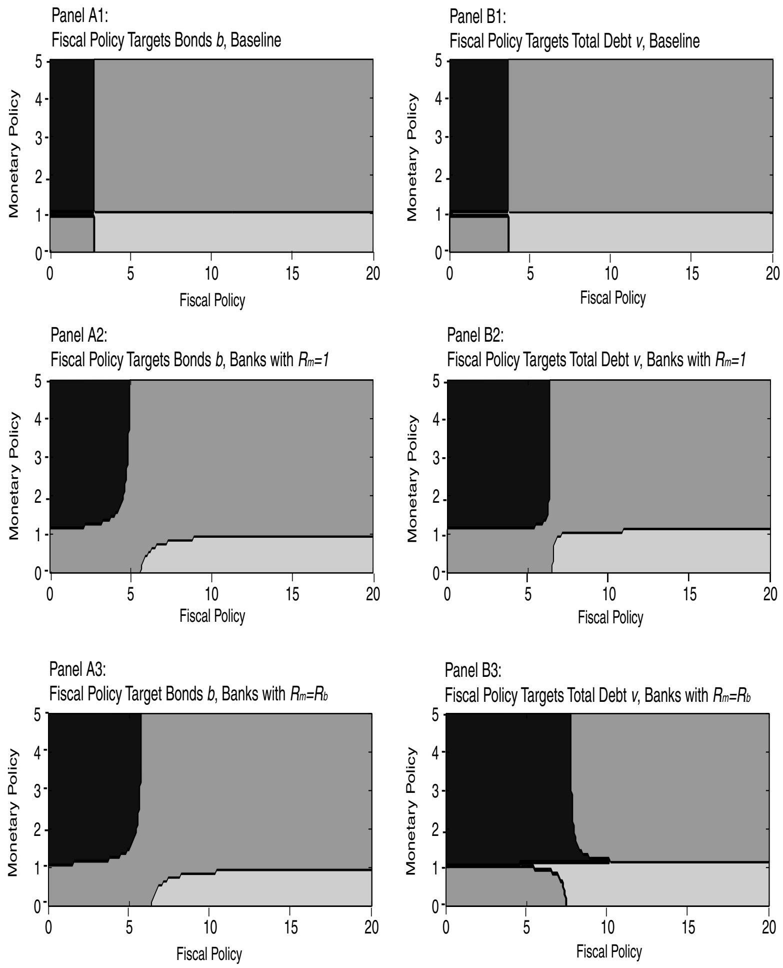
marginal utility of real balances and debt has to be zero. Thus, real balances and debt cannot grow too fast relative to the time discount factor.

[10] For a constant consumption path, $c_t = c$, and given policy targets for inflation and the debt-consumption ratio, equations (1)–(5) define a unique time-invariant solution for the endogenous variables, $x_t = x$, the steady state. I define a local approximation to the equilibrium in terms of small deviations from the steady state, which transforms the dynamic system of equations into a linear difference equation system. For a description of conditions for the existence and uniqueness of a bounded solution to linear difference equation system see, e.g., Sims (2000).

[11] Indeterminacy or nonexistence of an equilibrium raises an issue as to how useful the proposed theory is for the analysis of monetary policy. After all, we are trying to explain a particular outcome for the economy. Indeterminacy can be resolved by refining the equilibrium concept. For example, we might assume that decisions are coordinated on an extraneous random variable that has no relevance for the feasibility of outcomes, a sunspot. This gives rise to fluctuations as a result of self-fullfilling expectations. If no equilibrium exists for certain combinations of monetary and fiscal policy then we might conclude that some policy rules are simply not feasible in the long run (Sargent and Wallace [1981]).

[12] Figure 4 is based on a parameterization of the economy described in Section 3.

Figure 4 Price Level Determinacy



i.e., lump sum taxes respond strongly to government debt, then iteration on the transition equation for government debt defined by the government budget constraint (3) defines a unique bounded path for government debt. Conversely, if fiscal policy is active, i.e., lump sum taxes respond weakly to debt, then the unique bounded solution for debt from the government budget constraint defines debt as the discounted present value of future lump sum taxes and seigniorage revenue from money creation. This in turn defines a time path for the price level and thus the inflation rate. The implied time path for inflation need not be the same as the unique time path for inflation implied by an active monetary policy. Thus, active monetary and fiscal policies are inconsistent with the existence of an equilibrium. But if monetary policy is passive, then the difference equation describing the dynamics of inflation is consistent with a continuum of bounded solutions for inflation, in particular the inflation rate implied by the government budget constraint. This case is therefore also known as the fiscal theory of the price level. Finally, if monetary and fiscal policy are both passive, then there exists a continuum of bounded solutions to the system of difference equations, that is, the equilibrium is indeterminate.

Since for positive interest rates there is a uniquely defined demand for real balances, one can think of the interest rate as being supported by open market operations that supply the amount of money that is demanded at the given interest rate, equation (1). If the demand for real balances is characterized by a "liquidity trap"—that is, the demand is flat at a zero interest rate—then open market operations do not affect the equilibrium outcome.

3. INTEREST ON RESERVES AND THE CONDUCT OF MONETARY POLICY

I now describe a simple endowment economy with a banking sector that generalizes the baseline model described in the previous section. In this model banks are required to hold reserves, and one can study if and how the conduct of monetary policy needs to be changed once market interest rates are paid on reserves. I will limit attention to the question of how the payment of interest on reserves affects price level determinacy, that is, existence and uniqueness of an equilibrium.

An Economy with a Banking Sector

Consider a representative agent with preferences over a cash good, c, a credit good, k, real balances, m_h, real demand deposits, d, and real government bonds, b_h. Including these financial assets in preferences introduces a wedge into the asset pricing equations because the assets pay a liquidity premium. There is also a generic asset, a, that does not provide any liquidity services. The demand deposits are offered by a competitive banking sector that uses reserves and government bonds to service the demand deposits. The banking sector also makes loans, l, to the representative agent that are used to finance purchases of the credit good. Fiscal policy affects the evolution of government debt. The environment is a simplified version of Canzoneri et al. (2008).

Household Demand for Assets

The representative agent's preferences are

$$\sum_{t=0} \beta^t \left\{\ln c_t + \gamma_k \ln k_t + \gamma_m \ln m_{h,t} + \gamma_d \ln d_t + \gamma_b \ln b_{h,t}\right\} \tag{6}$$

and the budget constraint is

$$\begin{aligned} & c_t + k_t + m_{ht} + d_t + b_{ht} + a_t - l_t + \tau_t \\ \leq \quad & y_t + \left[m_{h,t-1} + d_{h,t-1}R_{dt} + b_{h,t-1}R_{bt} + a_{t-1}R_t - l_{t-1}R_{l,t}\right]/\pi_t, \end{aligned} \quad (7)$$

where the nominal interest rate for asset $j = m, d, b$, and l is R_j, the nominal interest rate on the generic asset is R, exogenous income is y, and lump sum taxes are τ. Real balances, demand deposits, and government bonds are assets that provide liquidity services in addition to being a store of value. The liquidity services are represented as direct contributions to a household's utility. The generic asset does not provide any liquidity services and is not included in the household's utility function. By assumption the household has to take out a loan to purchase the credit good

$$k_t \leq l_t. \quad (8)$$

The optimal choices of the household imply the following asset demand equations:

$$m_{ht} = \gamma_m \frac{R_{t+1}}{R_{t+1} - 1} c_t, \quad (9)$$

$$d_t = \gamma_d \frac{R_{t+1}}{R_{t+1} - R_{d,t+1}} c_t, \quad (10)$$

$$b_{ht} = \gamma_b \frac{R_{t+1}}{R_{t+1} - R_{b,t+1}} c_t, \quad (11)$$

$$l_t = \gamma_k \frac{R_{t+1}}{R_{t+1} - R_{l,t+1}} c_t. \quad (12)$$

Note that the household's demand for real balances is well-defined even at a zero nominal bond rate. The household's demand for real balances depends on the interest rate of the generic asset and not the bond rate. Furthermore, since bonds provide liquidity services, the bond rate will always be below the generic asset rate. Thus, even if the bond rate is zero the household demand for real balances is uniquely defined. There is no liquidity trap for household demand of real balances.

Intertemporal optimization with respect to the generic financial asset implies the Euler equation

$$1 = \left[\beta \frac{c_t}{c_{t+1}}\right] \frac{R_t}{\pi_{t+1}}, \quad (13)$$

where the term in square brackets is the marginal rate of substitution between consumption today and tomorrow. In the endowment economy equilibrium consumption of the cash and credit good is exogenous. With exogenous consumption, this Euler equation determines inflation conditional on the nominal interest rate for the generic asset.

Two remarks are in order. First, I deviate from the standard asset pricing setup to get potentially well-specified demand functions for real balances and demand deposits. Putting the assets into the utility function is one way to get well-defined demand functions. Alternatively, I could have assumed that these assets lower transactions costs and introduced the relevant cost terms in the budget constraint as in Goodfriend and McCallum (2007). Second, I want to have a simple model of bank lending, so just assume that the "credit" good has to be purchased through a one-period loan taken out from the bank.

Bank Demand and Supply of Assets

A bank takes in demand deposits that provide transactions services for the household and represent a liability to the bank. The bank's assets consist of loans made to the household, and bond and reserve holdings, b_b and m_b. The balance sheet of a bank is

$$l_t + b_{bt} + m_{bt} = d_t. \tag{14}$$

Banks need to hold reserves and bonds to service demand deposits:

$$b_{bt} + m_{bt} = \varphi d_t. \tag{15}$$

This equation represents an assumption on the bank's technology, namely what and how many assets the bank needs in order to generate the demand deposit services for the household. I assume that the bank uses liquid assets, i.e., bonds and reserves, in order to service demand deposits, but it need not hold 100 percent liquid assets, $\varphi < 1$. Furthermore, bonds and reserves are perfect substitutes in the production of demand deposit services.

Banks may also be forced to satisfy a reserve requirement that is imposed by a government regulator:

$$m_{bt} \geq \rho d_t. \tag{16}$$

Alternatively, the reserve ratio can reflect special precautionary preferences of banks for reserves. I assume that $\rho < \varphi$, otherwise banks would not hold other liquid assets besides reserves.[13]

I can assume that there is a representative bank that behaves competitively since the banking technology described above is characterized by constant returns to scale. Whereas banks receive interest on their bond holdings, the payment of interest on reserves (IOR), $R_m \geq 1$, is a policy choice. If bonds pay interest at a higher rate than do reserves, $R_b > R_m \geq 1$, then banks would prefer to hold bonds only against their demand deposits, but they are forced to hold at least a fraction, ρ, of their demand deposits in the form of reserves. If IOR is paid, I assume that interest is paid at the bond rate such that banks are

[13] Canzoneri et al. (2008) provide a more elaborate model of a banking sector that uses resources and not just assets to service demand deposits, and they allow for imperfect substitution between reserves and government bonds.

indifferent between reserves and bond holdings, $R_m = R_b$.[14] To summarize, the bank demand for reserves and bonds is determined by interest rates and reserve requirements as follows

$$m_{bt} = \rho d_t \text{ if } R_{b,t+1} > R_{m,t+1} \geq 1, \tag{17}$$

$$m_{bt} \in [\rho d_t, \varphi d_t] \text{ if } R_{b,t+1} = R_{m,t+1} \text{ or } R_{b,t+1} = 1, \tag{18}$$

$$b_{bt} = \varphi d_t - m_{bt}. \tag{19}$$

In any case, the zero profit condition for making loans and demand deposits determines the deposit rate

$$R_{d,t+1} = (1 - \varphi) R_{l,t+1} + (\varphi - \rho) R_{b,t+1} + \rho R_{m,t+1}. \tag{20}$$

This model for banks' reserve demand exhibits features of a "liquidity trap." First, at a zero bond rate the demand for reserves is indeterminate. Note, however, that the range of indeterminacy is bounded by the required reserve ratio and the desired liquid asset ratio. Second, once IOR is paid at the bond rate, the demand for reserves becomes indeterminate even at positive bond rates. Even though the banks' demand for reserves may be indeterminate within a range, the banks' joint demand for reserves and bonds is always uniquely determined.

Does the proposed "banking" technology make sense? For commercial banks the ratio of cash (including reserves with the Federal Reserve System) plus Treasury holdings relative to deposits has been remarkably stable from 1973 to the end of the 1980s (Figure 5). There was a sharp increase in the early 1990s and then a downward trend that has been reversed since last fall. At the same time, there was a steady decline of the ratio of cash relative to total deposits. Since excess reserves were small relative to required reserves before 2008, this must reflect a steady decline in the required reserve ratio.

Simultaneously with the introduction of IOR in the fall of 2008 and associated with various credit and liquidity programs, the amount of reserves banks hold with the Federal Reserve System has increased dramatically. These higher reserve holdings have not been accompanied by a corresponding decline of other liquid assets, such as treasuries or MBS, or by an increase of demand deposits (Figure 5). In terms of the proposed simple model this would have to be interpreted as a substantial increase in the desired ratio of liquid assets to deposits, φ.

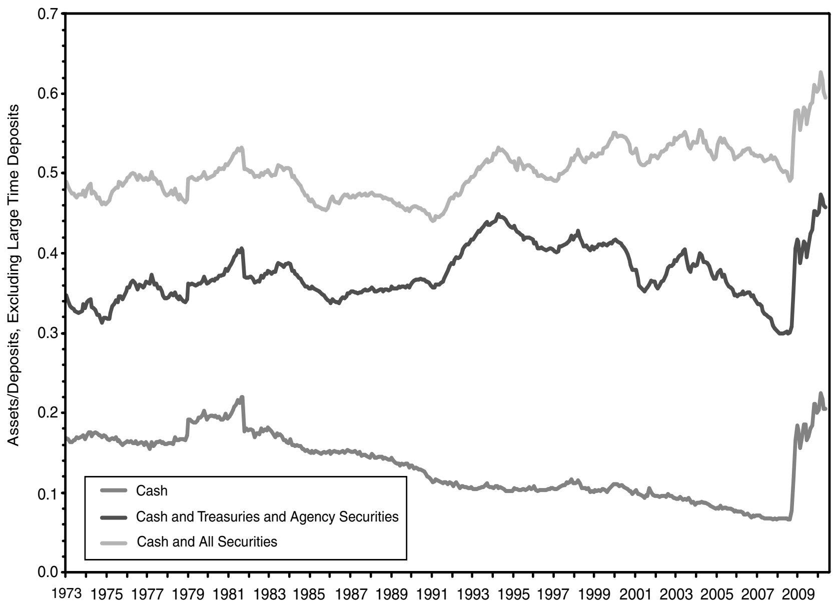
Government Supply of Assets

The government budget constraint is

$$b_t + m_t = \left[R_{b,t} b_{t-1} + m_{h,t-1} + R_{m,t} m_{b,t-1}\right] / \pi_t - \tau_t, \tag{21}$$

[14] In principle the policymaker could decide to make IOR greater than the bond rate, $R_m > R_b$, and reserves would dominate bonds as an asset for banks. I do not consider this case.

Figure 5 Liquid Asset Shares



Notes: The data are described in the Appendix.

where $b = b_h + b_b$ is the total amount of government bonds issued and $m = m_h + m_b$ is the monetary base. In an equilibrium the total amount of government debt has to equal the sum of bank and household bond holdings, and the monetary base has to equal the sum of bank reserves and household cash holdings.

Simplifying the Model

It is possible to simplify the exposition of the model considerably.[15] First, given the exogenous endowment of the cash and credit good, I can use the household demand for loans, (12), and the zero profit condition for banks, (20), to get an expression for the deposit rate:

$$R_{d,t+1} = \mathcal{R}_d \left(R_{t+1}, R_{b,t+1}, R_{m,t+1} \right). \tag{22}$$

[15] For the detailed derivation, see Hornstein (2010).

I can use this function in the household's demand for deposits, (10), and obtain the banks' demand for reserves:

$$\begin{aligned} m_{bt} &= \rho \mathcal{D}\left(R_{t+1}, R_{b,t+1}, R_{m,t+1}\right) c_t \text{ if } R_{b,t+1} > R_{m,t+1} = 1, \quad (23) \\ m_{bt} &\in [\rho, \varphi] \mathcal{D}\left(R_{t+1}, R_{b,t+1}, R_{m,t+1}\right) c_t \text{ if } R_{b,t+1} = R_{m,t+1} \text{ or } R_{b,t+1} = 1. \end{aligned}$$

Aggregate demand for real balances, the monetary base, is then the sum of household demand (9) for cash and bank demand for reserves (23):

$$m_t = \mathcal{M}\left(R_{t+1}, R_{b,t+1}, R_{m,t+1}\right) c_t. \quad (24)$$

The demand for monetary base inherits a "flat" indeterminacy range from the banks' reserve demand if the bond rate is zero or interest is paid on reserves.

Analogously to the total demand for real balances, I can define a total demand for government bonds by households and banks:

$$b_t = \mathcal{B}\left(R_{t+1}, R_{b,t+1}, R_{m,t+1}\right) c_t. \quad (25)$$

Corresponding to the aggregate demand for real balances, the aggregate demand for bonds also inherits a "flat" indeterminacy range from the banks' demand for bonds. Aggregate demand for total government debt is the sum of the demand for real balances and bonds, equations (24) and (25),

$$v_t = \mathcal{V}\left(R_{t+1}, R_{b,t+1}, R_{m,t+1}\right) c_t. \quad (26)$$

As pointed out above, banks' demand for reserves and bonds together is always determinate and the same then applies to the demand for total government debt (money and bonds).

The reduced form of the economy can now be represented by the following set of equations:

$$\begin{aligned} m_t &= \mathcal{M}\left(R_{t+1}, R_{b,t+1}, R_{m,t+1}\right) c_t, && (27) \\ 1 &= \beta \frac{c_t}{c_{t+1}} \frac{R_{t+1}}{\pi_{t+1}}, && (28) \\ v_t &= \frac{R_{b,t} v_{t-1} - \left(R_{b,t} - 1\right) \tilde{m}_{t-1}}{\pi_t} - \tau_t, && (29) \\ \tilde{m}_t &= \tilde{\mathcal{M}}\left(R_{t+1}, R_{b,t+1}\right) c_t, && (30) \\ v_t &= \mathcal{V}\left(R_{t+1}, R_{b,t+1}, R_{m,t+1}\right) c_t, && (31) \\ v_t &= b_t + m_t. && (32) \end{aligned}$$

Equation (27) is the aggregate demand for real balances. Equation (28) is the household Euler equation for the generic asset, (13). Equation (29) is the government budget constraint in terms of total debt outstanding v, and $\tilde{m}$ denotes non-interest-bearing government debt. Without interest on reserves, non-interest-bearing debt is aggregate real balances, $\tilde{m} = m$; and with interest on reserves, non-interest-bearing debt is cash holdings by households, $\tilde{m} = m_h$. Equation (31) is the aggregate demand for government debt. Equation

(32) defines total government debt as the sum of real balances and bonds. The baseline model, (1)–(4), is obtained from Section 2 if one assumes that bonds and demand deposits do not provide any liquidity services, $\gamma_b = \gamma_d = 0$; eliminates the credit good, $\gamma_k = k = 0$; and eliminates the banking sector.

Price Level Determinacy with Interest on Reserves

I now show that the simple baseline model from Section 2 and the just described model with a banking sector have very similar implications for how monetary and fiscal policy affect price level determinacy. Whether or not interest is paid on reserves, the model with banking does not materially affect this result. In particular, it appears that the volume of bank reserves does not matter.

The reduced form representation of the economy with a banking sector, equations (27)–(32), appears to be slightly more complicated than the simple baseline model, equations (1)–(4), but the structure of the two economies is very similar. In order to close the model with banking, I again assume that there are fixed endowments of the consumption good, cash and credit; and specify monetary and fiscal policy as responding to inflation and government debt, equation (5). I again study the local properties of the linearized dynamic system defined by equations (27)–(32) and the policy rules (5). In the baseline model, fiscal policy responds to the stock of outstanding real bonds, b, that is, interest-bearing government debt. For reasons that will immediately become apparent, I also consider a fiscal policy that responds to the total stock of government debt, v. I can also do that for the simple baseline model and, comparing Panels A1 and B1 of Figure 4, it is clear that this has no substantial impact on the issue of equilibrium existence and uniqueness.

In order to characterize the implications of monetary and fiscal policy for price level determinacy I need to parameterize the model with banking. Relative to the baseline model, I need to make assumptions on households' steady-state asset holdings (real balances, m , bonds, b, and deposits, d); banks' required reserve ratio, ρ, and desired liquidity, φ; and on steady-state rates of return on the generic asset, R, bonds, R_b, and money, π. I follow Canzoneri et al.'s (2008) calibration of the 1990–2005 U.S. economy. The time period is assumed to be a quarter. The household steady-state ratios of real balances, bonds, and demand deposits to consumption are $m_h/c = 0.3$, $b_h/c = 0.9$, and $d/c = 2.45$. Steady-state nominal interest rates on reserves, bonds, and the generic asset are $R_m = 1$, $R_b = 1.011$, and $R_a = 1.015$. Steady-state inflation is $\pi = 1.007$. The reserve ratio is $\rho = 0.05$ and reflects the ratio of vault cash and bank deposits with the Federal Reserve. The desired liquidity ratio is $\varphi = 0.30$ and reflects the ratio of bank holdings of treasury debt, agency debt, agency MBS, and total reserves to total deposits.

Fiscal Policy Targets Total Debt

Suppose first that fiscal policy targets total debt, v, and that no interest is paid on reserves. Comparing Panels B1 and B2 of Figure 4 it is apparent that the parameter regions that characterize equilibrium existence and uniqueness are qualitatively similar to the baseline model. Price level determinacy is obtained in the northeast region (active monetary policy and passive fiscal policy) and the southwest region (passive monetary policy and active fiscal policy) of the parameter space.

Now suppose that fiscal policy continues to target total debt, but interest is paid on reserves. Because of interest on reserves, total demand for real balances is indeterminate for a range that depends on the reserve ratio and the desired liquidity ratio of banks. Even if the total supply of real balances falls into that range, this does not create a problem for the conduct of monetary policy.

Consider equations (28)–(31) of the reduced form together with the monetary and fiscal policy rules. These equations are sufficient to determine an equilibrium in terms of the inflation rate, interest rates, and total debt, $\left\{\pi_t, R_{t+1}, R_{b,t+1}, v_t\right\}$, if the equilibrium exists. The allocation of total government debt between interest-bearing reserves and interest-bearing debt is irrelevant. In particular, the magnitude of reserves at banks does not matter, as long as the reserves remain within the range of indeterminacy.

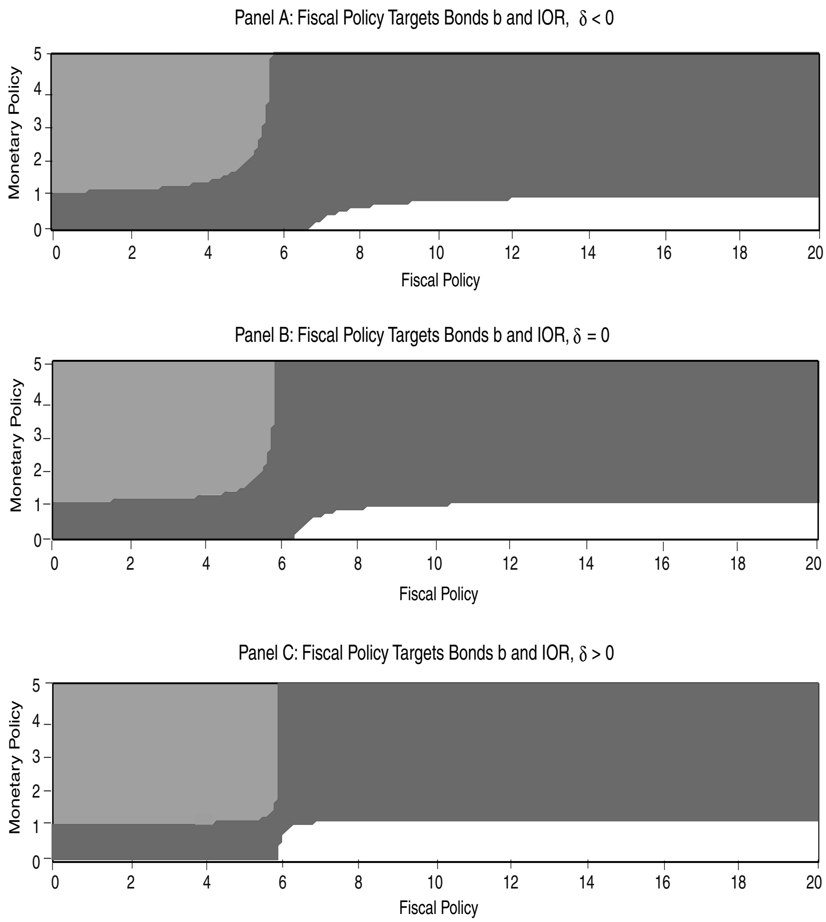
Comparing Panels B2 and B3 of Figure 4 shows that paying interest on reserves has some impact on the issue of price level determinacy. If there is price level determinacy in the northeast region of the parameter space without IOR, then for a given active monetary policy, fiscal policy with IOR has to be somewhat more passive in order for the equilibrium to remain unique.[16] Conversely, in the southwest region of the parameter space, for a given monetary policy, fiscal policy with IOR needs to be more active to obtain price level determinacy.

Fiscal Policy Targets Real Bonds

Now suppose that fiscal policy targets the stock of real bonds, b, rather than total debt, v, but no interest is paid on reserves. Comparing Panels A2 and B2 of Figure 4 shows that for any given monetary policy, fiscal policy can be somewhat more active before losing price level determinacy, either because of nonexistence or nonuniqueness. But now it appears that there is a problem if interest is paid on reserves, since the demand for government bonds—

[16] Recent projections of rapidly expanding fiscal deficits might suggest that fiscal policy has shifted toward a more active stance, that is, taxes are responding less strongly to outstanding debt. If monetary policy were to remain active, fiscal and monetary policy could become inconsistent, that is, an equilibrium would not exist. Thus, the payment of interest on reserves might require a further adjustment of either monetary or fiscal policy to maintain the existence of an equilibrium.

Figure 6 Monetary Policy Targets Reserves



and relatedly the demand for real balances—becomes indeterminate for some range. A well-defined demand for government bonds is, however, needed, since fiscal policy is supposed to respond to the stock of outstanding bonds.

I can resolve the indeterminacy of the demand for bonds through the introduction of an additional policy rule that determines their equilibrium values. For example, the central bank might conduct open market operations (OMO) that adjust real balances in response to the inflation rate:

$$m_t = h(\pi_t), \tag{33}$$

with an elasticity of δ. In other words, because the money demand equation, (27), no longer determines real balances, monetary policy can choose real balances.[17]

Figure 4, Panel A3 graphs the parameter regions for price level determinacy when monetary policy does not adjust real balances in response to the inflation rate, $\delta = 0$. The impact of paying interest on reserves relative to not paying interest on reserves, Figure 4, Panel A2, is similar to the case when fiscal policy targets total debt and not bonds only.

How much paying IOR matters now also depends on the new OMO parameter, δ. Figure 6 displays the parameter regions for price level determinacy when fiscal policy targets real bonds and the OMO parameters are $\delta = 100$ (Panel A), $\delta = 0$ (Panel B), and $\delta = -100$ (Panel C). Given that the OMO response to real balances is essentially a response to bank reserves, one might think that with IOR, monetary policy would have to target both inflation and bank reserves. This interpretation has to be qualified for two reasons. First, bank reserves matter only because I have assumed that fiscal policy targets bonds and not total debt. Second, the graphs in Figure 6 are based on very extreme values for the OMO policy parameter. For δ values that are of similar magnitude as the monetary and fiscal parameters, α and γ, the parameter regions for price level determinacy are essentially the same.

4. CONCLUSION

This article addresses the question of whether paying interest on the reserve accounts that banks hold with a central bank affects the conduct of monetary policy. For this purpose I introduce a stylized model of banks that hold reserves into a standard baseline model of money. This model suggests that paying interest on reserves does not drastically change the implications of monetary policy, implemented as an interest rate policy, for price level determinacy. Furthermore, the amount of outstanding reserves does not appear to be critical for issues of price level determinacy.

The scope of the article is rather narrow. For example, I do not study how the payment of IOR affects the dynamic response of the economy to shocks for given monetary and fiscal policy rules. The model can be used to address this issue if features are added that make money non-neutral, for example, a New Keynesian Phillips curve based on sticky prices. Preliminary results for such an augmented model suggest that for the same monetary and fiscal policy rules the dynamic response of inflation and output to shocks does

[17] We usually think of OMO as determining nominal quantities. I have chosen a policy rule that chooses real balances to keep the exposition simple. One could interpret the proposed policy rule as responding to inflation and to the price level. Alternatively, one could simply start with a policy rule that sets the nominal money stock and study the more complicated system.

depend on whether or not interest is paid on reserves, but the differences are not substantial.

The effects of financial market interventions by central banks, however, cannot be studied in this framework. Since the model's concept of liquidity for the financial sector is rather narrow, the model has nothing to say about central bank provision of liquidity to banks through an increase of the banks' reserve accounts. For example, the model does not provide any rationale for the Federal Reserve's program to purchase agency MBS as opposed to other government debt. Indeed, the simple banking model assumes that agency MBS and treasuries provide the same liquidity services to banks.[18] For a critical review of the Federal Reserve interventions in specific financial markets that gave rise to the expansion of the Federal Reserve's balance sheet, in particular, the volume of reserve liabilities, see Hamilton (2009).

APPENDIX

Figure 2 displays daily data for the federal funds target set by the FOMC and the effective federal funds rate from January 2000–February 2010. In addition, Panel C of Figure 2 also displays the interest rate that was paid on required reserves and on excess reserves from September 2008 on. Figure 3 displays monthly averages from January 1980–February 2010 for the following short-term interest rates: the effective federal funds rate, the three-month constant maturity Treasury rate, the three-month nonfinancial commercial paper rate, the rate for three-month certificates of deposit in the secondary market, and the prime bank lending rate. Figure 5 displays monthly liquid asset ratios of all commercial banks, domestically chartered and foreign related institutions, from January 1973–January 2010 based on the Federal Reserve Board's H.8 table. Securities in bank credit include Treasury and agency securities and other securities. A large part of agency securities consists of MBS issued by GSEs such as the Government Mortgage Association (Ginnie Mae, GNMA), the Federal National Mortgage Association (Fannie Mae, FNMA), or the Federal Home Loan Mortgage Corporation (Freddie Mac, FHLMC). Other securities include private label MBS, among others. Cash includes vault cash and reserves with the Federal Reserve. The liquid asset ratio is calculated relative to bank deposits excluding large time deposits. All series are from Haver.

[18] Given that the GSEs Fannie Mae and Freddie Mac have become wards of the federal government, this does not appear to be such an unreasonable assumption. Indeed, the only reason to distinguish between Treasury debt on the one hand and agency-issued debt and MBS on the other hand appears to be political: GSE-issued debt does not count toward the congressionally mandated federal debt limit.

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Are Wages Rigid Over the Business Cycle?

Marianna Kudlyak

The search and matching model of the labor market has become a leading model of unemployment in macroeconomics. Recent work (Shimer 2005) shows that under common parameter values the standard search and matching model cannot account for the cyclical volatilities of vacancies and unemployment observed in the data.[1] This difficulty is related to the flexibility or, alternatively, rigidity of real wages over the business cycle.[2] In this article, I review empirical evidence on wage flexibility as it relates to search and matching models.

The search and matching model introduces frictions into the labor market in the sense that workers and employers cannot costlessly contact each other. In an economy with frictions, market prices are not competitively determined, and the standard search and matching framework assumes that wages are determined by a particular solution to a bargaining problem between workers and employers, the Nash bargaining solution. Under this bargaining, wages increase when productivity is high, thus limiting incentives for job creation. Hall (2005) and Shimer (2005) show that replacing the Nash bargaining solution with an alternative wage determination procedure that makes wages more rigid amplifies the volatility of unemployment and vacancies the model can generate.

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[1] The "standard search and matching model" refers to the model studied by Shimer (2005) and developed in Mortensen and Pissarides (1994). Pissarides (2000) provides a textbook exposition of the standard search and matching model.

[2] The terms "rigid" and "acyclical" are used interchangeably and imply a lack of response to a cyclical indicator.

This line of research motivates important questions: How flexible, or rigid, are real wages over the business cycle? Is observed wage rigidity consistent with the wage rigidity needed to amplify the fluctuations in the textbook search and matching model?

In the job creation decision, a firm takes into account not only the initial wage in a newly formed match but also the entire expected stream of future wages to be paid in the match. Thus, job creation in frictional labor markets places the focus on the cyclical behavior of the expected present discounted value of wages. The volatility of unemployment in the model depends on the intensity of job creation through changes in the job finding rate. As Shimer (2004) emphasizes, what is relevant for the volatility of job creation, and, thus, of unemployment, is the rigidity of the present discounted value of wages that, at the time of hiring, a firm expects to pay to a worker over the course of the employment relationship.

A large empirical literature exists that studies the behavior of individual wages over the business cycle. The literature finds that the wages of newly hired workers are more cyclical than wages of workers in ongoing employment relationships (for example, Bils [1985] and Carneiro, Guimarães, and Portugal [2009]).

One crucial aspect of the existing empirical literature is that it provides evidence on the cyclical behavior of the current wage, but not on the cyclical behavior of the expected present discounted value of future wages within a match formed in the current period. Emphasizing the importance of the future wages to be paid in the long-term employment relationships, Kudlyak (2007) estimates the cyclicality of the user cost of labor, which is the difference between the expected present discounted value of wages paid to a worker hired in the current period and the value paid to a worker hired the following period. Kudlyak constructs the user cost of labor from the individual wage and turnover data. She finds that the user cost of labor is more cyclical than wages of newly hired workers. Haefke, Sonntag, and van Rens (2009) argue for the importance of the elasticity of the expected present discounted value of wages with respect to the expected present discounted value of productivity in newly formed matches, which they refer to as permanent values of wages and productivity, respectively. They do not estimate the elasticity directly but, using model simulations, conclude that the elasticity of the current period wage of newly hired workers with respect to current period productivity "constitutes a good proxy for the elasticity of the permanent wage with respect to permanent productivity for the case of instantaneously rebargained wages" and "can be seen as a lower bound for [the elasticity of the permanent wage with respect to permanent productivity] in the case of wage rigidity on the job."

Pissarides (2009), Haefke, Sonntag, and van Rens (2009), and Kudlyak (2009) study whether the search and matching model can simultaneously match the empirical wage and unemployment statistics. Pissarides (2009) and Haefke, Sonntag, and van Rens (2009) compare the elasticity of wages in the model to the elasticity of wages of newly hired workers in the data. They conclude that making wages in the model rigid enough to generate the observed volatility of unemployment requires more rigid wages than are found in the data. Kudlyak (2009) shows that the free entry condition in the model ties together match productivity and the user cost of labor. She calibrates the model to match the estimated cyclicality of the user cost of labor and concludes that the model calibrated to wage data cannot generate the empirical volatilities of the vacancy-unemployment ratio.

In summary, in the model, the rigid expected present discounted value of wages in newly formed matches amplifies the response of firm's surplus to productivity shocks. This increases the volatility of job creation and thus of unemployment. The success of the model in generating the empirical volatilities of vacancies and unemployment depends on whether the required rigidity of the relevant measure of wages is consistent with the data. The studies reviewed suggest that wages in the data may not be as rigid as required for generating empirical volatilities of vacancies and unemployment in the standard search and matching model.

The remainder of the paper is structured as follows. Section 1 summarizes the textbook search and matching model and the unemployment volatility puzzle. Using an example, I demonstrate the importance of the expected present discount value of wages for the job creation decision. Section 2 surveys empirical evidence on the cyclicality of individual wages of the newly hired workers and wages of workers in ongoing relationships. Then I review empirical evidence on the cyclicality of a measure of wages that takes into account the expected present discounted value of future wages. Section 3 concludes.

1. SEARCH AND MATCHING MODEL

The Standard Model

An economy is populated by a continuum of firms and a continuum of measure 1 workers. Both firms and workers are risk neutral and infinitely lived. Firms maximize the present discounted value of profits. Workers maximize the present discounted value of wages and do not value leisure. Firms and workers discount the future with a common discount factor β, $0 < \beta < 1$. Time is discrete.

A firm can choose to remain inactive or to start production, which requires only labor input. To start production, a firm must enter the labor market and hire a worker. Upon entering the labor market, a firm opens a vacancy and searches for a worker. During each period a firm must pay a per vacancy cost, c. An unemployed worker receives a per period unemployment compensation, b, and costlessly searches for a job. Employed workers earn wages and cannot search. When a firm with an open vacancy and an unemployed worker meet, they form a match that immediately becomes productive. While matched, all firm-worker pairs have the same constant return to scale production technology, which uses a unit of labor indivisibly supplied by the worker. Each firm-worker match produces per period output z, thus z is also aggregate productivity. z evolves stochastically according to the first-order Markov process. Every period, a firm-worker pair separates exogenously with probability δ.

Given the number of unemployed workers, u, and the number of vacancies, v, the number of newly created matches in the economy is determined by a matching function, $m(u, v)$. It is assumed that $m(u, v) = Ku^{\alpha}v^{1-\alpha}$, where $0 < \alpha < 1$ (Petrongolo and Pissarides 2001) and K is a positive constant. Let θ denote labor market tightness, i.e., $\theta \equiv \frac{v}{u}$. Let $q(u, v) \equiv \frac{m(u,v)}{v} = K\theta^{-\alpha}$ denote the probability of filling a vacancy for a firm. Let $\mu(u, v) \equiv \frac{m(u,v)}{u} = K\theta^{1-\alpha}$ denote the probability of finding a job for an unemployed worker. Thus, the unemployment in this economy evolves according to the following equation, given u_0:

$$u_{t+1} = u_t + (1 - u_t)\,\delta - \mu\,(u_t, v_t)\,u_t.$$

Dropping the time subscripts and using $'$ to denote the next period values, I summarize the value functions of a worker and of a firm as follows. The value function of a firm with a worker is

$$J\,(z) = z - w\,(z) + \beta\,(1 - \delta)\,E_z J\left(z'\right), \tag{1}$$

where z' denotes productivity in the next period and E_z is a conditional expectations operator. Equation (1) takes into account that in each period with probability δ, the firm-worker match separates and the firm obtains a value of an inactive firm, which is 0. The value function of an opened vacancy is

$$V\,(z) = -c + q\,(\theta\,(z))\,J\,(z) + (1 - q\,(\theta\,(z)))\,\beta E_z V\left(z'\right). \tag{2}$$

The value function of an employed worker is

$$W\,(z) = w\,(z) + \beta\,(1 - \delta)\,E_z W\left(z'\right) + \beta\delta E_z U\left(z'\right). \tag{3}$$

The value function of an unemployed worker is

$$U(z) = b + \beta E_z\left(\mu(\theta(z'))W(z') + (1 - \mu(\theta(z')))U(z')\right). \tag{4}$$

There are two important conditions in the standard model. First, there is free entry for firms, i.e., firms enter the labor market and post vacancies until the value of an open vacancy, $V(z)$, equals the value of an inactive firm, 0. From (2) free entry implies the following condition:

$$\frac{c}{q\left(\theta\left(z\right)\right)} = J\left(z\right). \tag{5}$$

Second, wages are rebargained every period in new and ongoing matches according to the Nash bargaining rule. The Nash bargaining rule yields the following division of the surplus from the match, $S(z) \equiv J(z) + W(z) - V(z) - U(z)$, in every period:

$$J(z) = (1-\eta)S(z), \tag{6}$$

$$W(z) - U(z) = \eta S(z), \tag{7}$$

where η is a worker's bargaining power, $0 < \eta < 1$. Equations (6)–(7) imply that each party obtains a constant share of the surplus from the match.

Using equations (1)–(5) yields the following equation for the surplus:

$$S(z) = z - b + \beta E_z((1-\delta) - K\theta(z')^{1-\alpha}\eta)S(z'). \tag{8}$$

Combining (5) and (6) yields the job creation condition

$$\frac{c}{K\theta\left(z\right)^{-\alpha}} = (1-\eta)\,S\left(z\right). \tag{9}$$

Combining (8) with the job creation condition yields the following equation for the evolution of the vacancy-unemployment ratio:

$$\frac{c}{K\theta(z)^{-\alpha}} = (z-b)(1-\eta) + \beta E_z((1-\delta) - K\theta(z')^{1-\alpha}\eta)\frac{c}{K\theta(z')^{-\alpha}}. \tag{10}$$

Equation (8) links the evolution of the vacancy-unemployment ratio, θ, to the productivity shock, z. Using this equation and common parameter values, Shimer (2005) shows that the standard search and matching model cannot generate the volatilities of vacancies and unemployment observed in the data. In particular, in the U.S. data, the standard deviation of the vacancy-unemployment ratio is 20 times larger than the standard deviation of labor productivity. The standard search and matching model predicts the volatility of the vacancy-unemployment ratio as almost one-to-one to the volatility of the productivity. Since productivity shocks are the driving force in the model, the model is said to lack an internal propagation mechanism.[3] This failure of the standard search and matching model to generate empirical volatilities of vacancies and unemployment is often referred to as the unemployment volatility puzzle (Pissarides 2009).

[3] See Hornstein, Krusell, and Violante (2005) for a detailed inspection of the mechanism. See also Costain and Reiter (2008).

Rigid Wages Within Matches

To understand what measure of wages affects allocations in the search and matching model, consider the following modification of the standard model. In the standard search and matching model, wages in both newly formed and ongoing matches are set in each period using the Nash bargaining rule. In the modified model, wages in newly formed matches are set using the Nash bargaining rule but wages in ongoing matches remain constant and equal to the wage at the time of hiring. We will see that the modified model delivers the same equation for the vacancy-unemployment ratio, thus the same allocations, given the initial conditions, as the standard model despite the fact that in the modified model wages in ongoing matches are rigid. The modified model is a discrete time version of the argument presented in Shimer (2004).[4]

In the standard model, given that all matches are equally productive, wages of new hires and existing workers are equal in each period. This implies that when the aggregate productivity is z_t, the value of a firm with a worker in an ongoing match that started in period t_0, $J^{t_0}(z_t)$, equals the value of a firm in the newly formed match, $J^t(z_t)$. Similarly, when the aggregate productivity is z_t, the value of an employed worker in an ongoing match that started in period t_0, $W^{t_0}(z_t)$, equals the value of a newly hired worker in t, $W^t(z_t)$. In the modified model, these values are not necessarily equal. Dropping the time subscripts, using z_0 to denote the aggregate productivity at the time a match is formed and using $'$ to denote the next period values, we can summarize the value functions in the modified model as follows:

$$
\begin{aligned}
J^{z_0}(z) &= z - w(z_0) + \beta(1-\delta) E_z J^{z_0}(z'); \\
V(z) &= -c + K\theta(z)^{-\alpha} J^z(z) + \left(1 - K\theta(z)^{-\alpha}\right) \beta E_z V(z'); \\
W^{z_0}(z) &= w(z_0) + \beta(1-\delta) E_z W^{z_0}(z') + \beta\delta E_z U(z'); \\
U(z) &= b + \beta E_z \left[K\theta(z')^{1-\alpha} W^{z'}(z') + \left(1 - K\theta(z')^{1-\alpha}\right) U(z') \right].
\end{aligned}
$$

In the modified model the free entry condition, (5), and the surplus division rule, (6)–(7), are required to hold only for the values at the time of hiring, which can be denoted as $J^z(z)$ for a firm and $W^z(z)$ for a worker. Thus, in the modified model, equations corresponding to equations (6), (7), and (9) are as follows:

$$J^z(z) = (1-\eta) S^z(z), \tag{11}$$

[4] See also Haefke, Sonntag, and van Rens (2009) and Pissarides (2009) for insightful discussions of this example and Rudanko (2009) for a model with endogeneous wage rigidity within ongoing matches.

$$W^z(z) - U(z) = \eta S^z(z), \tag{12}$$

$$\frac{c}{q(\theta(z))} = (1-\eta)S^z(z), \tag{13}$$

where $S^z(z)$ is the surplus from a newly formed match when the aggregate productivity is z, $S^z(z) \equiv J^z(z) + W^z(z) - V(z) - U(z)$.

To demonstrate that this modified model delivers exactly the same allocations as the standard model (in which wages are rebargained in all matches every period), it suffices to show that it delivers the same equation for the vacancy-unemployment ratio as the standard model, (10). Using equations for $J^{z_0}(z)$,$W^{z_0}(z)$, and $U(z)$, one can derive the following equation for the total surplus of the newly formed match:

$$S^z(z) = z - b + \beta(1-\delta)E_z S^z(z') + \beta\eta E_z \mu(\theta(z'))S^{z'}(z'), \tag{14}$$

where z' is productivity in the next period.

Note, however, that $J^z(z') = J^{z'}(z') + \left(\frac{1}{1-\beta(1-\delta)}\right)(w(z') - w(z))$, where $\left(\frac{1}{1-\beta(1-\delta)}\right) w(z)$ is a present discounted value of wages paid to a worker hired when the aggregate productivity is z. Similarly, $W^z(z') = W^{z'}(z') - \left(\frac{1}{1-\beta(1-\delta)}\right)(w(z') - w(z))$. Then $S^z(z') = J^z(z') + W^z(z') - U(z') = J^{z'}(z') + W^{z'}(z') - U(z') = S^{z'}(z')$. Substituting $S^z(z')$ for $S^{z'}(z')$ in (14) yields

$$S^z(z) = z - b + \beta E_z((1-\delta) - \mu(\theta(z'))\eta)S^{z'}(z'), \tag{15}$$

where $S^z(z)$ is the surplus from a newly created match when the aggregate productivity is z and $S^{z'}(z')$ is the surplus from a newly created match when the aggregate productivity is z'.

Substituting $S^z(z)$ from the job creation condition (13) into (15) yields exactly the same equation for the vacancy-unemployment ratio, θ, as in the standard model, equation (10):

$$\frac{c}{q(\theta(z))} = (z-b)(1-\eta) + \beta E_z((1-\delta) - \mu(\theta(z'))\eta)\frac{c}{q(\theta(z'))}.$$

The two models considered above have important similarities. Both models deliver the same total surplus at the time of hiring, and it is split between a worker and a firm by the same rule. However, they differ in how the wages are determined within ongoing employment relationships. In the standard model, wages are renegotiated for every match in every period. Because all matches are equally productive, this implies that newly hired workers and workers in existing employment relationships receive the same wages. In the modified model, wages of newly hired workers are determined based on the aggregate productivity at the time of hiring. Once set, the wage remains rigid throughout the duration of the match. The modified model generates wages of newly hired workers that are more responsive to the aggregate conditions than wages of the workers in ongoing relationships. However, in the modified model, the rigidity of wages within employment relationships does not affect allocations. Thus, this example shows that the rigidity of wages in ongoing matches is not sufficient to amplify the volatility of the vacancy-unemployment ratio.

To understand what kind of wage rigidity has an effect on allocations in the model, rewrite the job creation (13) using $(1-\eta)S^z(z) = J^z(z)$ and using the expression for $J^z(z)$ in the sequential form. This yields:

$$\frac{c}{q(\theta(z_t))} = E_t \sum_{\tau=t}^{\infty} (\beta(1-\delta))^{\tau-t} z_\tau - E_t \sum_{\tau=t}^{\infty} (\beta(1-\delta))^{\tau-t} w_{t,\tau}(z_t), \quad (16)$$

where $w_{t,\tau}$ is a period τ wage of a worker hired in period t. Equation (16) shows the relationship between the labor market tightness, $\theta(z_t)$, and the expected present discounted value of wages, $E_t \sum_{\tau=t}^{\infty} (\beta(1-\delta))^{\tau-t} w_{t,\tau}(z_t)$, given productivity z_t. Note that $E_t \sum_{\tau=t}^{\infty} (\beta(1-\delta))^{\tau-t} z_\tau$ is a function of z_t alone. Both $\theta(z_t)$ and $E_t \sum_{\tau=t}^{\infty} (\beta(1-\delta))^{\tau-t} w_{t,\tau}(z_t)$ change in response to changes in z_t. The extent of the response of $\theta(z_t)$ to z_t depends on the extent of the response of the expected present discounted value of wages to be paid in a new employment relationship that starts at t. However, it does not depend on the change of wages within the employment relationship if this change does not affect the expected present discounted value of wages to be paid in a new match.

2. EMPIRICAL EVIDENCE ON CYCLICAL BEHAVIOR OF WAGES

This section reviews the empirical evidence on wage cyclicality. First, I present the empirical evidence on the behavior of individual wages over the business cycle, distinguishing wages of new hires from wages of workers in ongoing employment relationships (often referred to as job stayers). Second, I present the empirical evidence on the history dependence of wages. Then, I present the evidence on the cyclical behavior of a measure of wages that takes into account both the initial wage and the expected value of future wages paid in the newly formed matches. Finally, I summarize the quantitative implications of the evidence for the volatility of vacancies and unemployment in the standard search and matching model.

Cyclicality of Wages of New Hires and Wages of Existing Workers

Below I provide a statistical model of individual wages with a particular emphasis on how wages depend on the unemployment rate. Then I survey results from the empirical studies that include information on individual workers and from the studies that include information on workers and their employers.[5] The findings show that wages of newly hired workers are more procyclical than wages of job stayers. All of these studies refer to current wages and not to the expected present discounted value of future wages.

General Framework

In labor economics the standard statistical model for wages is Mincer regression, which attributes variation in the logarithm of wages to the observable characteristics of a worker—years of schooling, a quadratic polynomial in labor market experience, and other factors (Mincer 1974). These variables are supposed to reflect productivity (or human capital) differences. The literature that studies the behavior of individual wages over the business cycle includes the contemporaneous unemployment rate as a business cycle indicator. What is of interest for the questions in this article are the differences, if any, of the responses of wages of workers in ongoing matches (job stayers) and wages of new hires to changes in the unemployment rate.

The individual wage equation that distinguishes between the cyclical response of wages of job stayers and wages of new hires is specified as follows:

$$\ln w_{it} = X_i \alpha + X_{it}\gamma + \beta U_t + \beta^{nh} U_t * I_{it}^{nh} + \delta I_{it}^{nh} + \eta_i + v_{it}, \qquad (17)$$

where w_{it} is a real wage of worker i in t, X_i is a vector of observable individual-specific explanatory variables that remain fixed over time, X_{it} is a vector of individual controls that vary with time, U_t is a measure of the unemployment rate, and η_i and v_{it} are the unobservable error terms. I_{it}^{nh} is a dummy variable that takes value 1 if an individual is a new hire, and 0 otherwise. The new hire is defined as a worker who has been employed at a firm for less than a specified period, usually one year. Error terms are assumed independent of each other and of all explanatory variables in X. The variables commonly included in X_{it} are a quadratic in worker labor market experience and a quadratic in tenure (for job stayers). Because of the structure of the survey data, the time period is typically one year.

[5] Given a large literature, this survey does not aim to summarize all works on the real wage cyclicality. An interested reader is referred to the surveys in Abraham and Haltiwanger (1995) and Brandolini (1995).

The coefficient on the unemployment rate is interpreted as a semi-elasticity, which indicates a percent change in wage in response to a one percentage point increase in the unemployment rate. If the semi-elasticity is positive, the wage is called procyclical, i.e., it moves positively with the business cycle. For job stayers the cyclicality is measured by β. If the cyclicality of wages of new hires differs from the cyclicality of job stayers, then the coefficient on the interaction term, β^{nh}, is statistically significantly different from zero and the cyclicality of new hires is measured by $\beta + \beta^{nh}$.

Evidence from Worker Survey Data

Most of the existing evidence on the cyclicality of individual wages comes from studies that use individual level survey data: the National Longitudinal Survey (NLS), the Panel Study of Income Dynamics (PSID), and the National Longitudinal Study of Youth (NLSY). These data allow tracking individual workers' histories across time and contain information on individual workers' characteristics, including education, age, sex, and job characteristics such as industry and occupation.[6]

Bils (1985) is the first study that examines the cyclicality of individual wages while separating the wages of job stayers from the wages of new hires. He also distinguishes between new hires who are hired from another job and new hires who are hired from unemployment. Using the individual data on men from NLS for the period 1966–1980, Bils finds that as the unemployment rate increases by one percentage point, individual wages of white male workers on average decrease by 1.59 percent. Once the job changers are explicitly accounted for, the results show that wages of job changers are much more procyclical than wages of job stayers. In particular, wages of job stayers decrease by 0.64 percent while wages of job changers decrease by 3.69 percent in response to one percentage point increase in the unemployment rate. Similarly, wages of workers who move in and out of employment are also more procyclical than wages of workers who do not change jobs.

Shin (1994), using a different estimation procedure for the NLS data on men's wages from 1966–1981, estimates separate equations for workers who remain with the same employer from $t - 1$ to t and for workers who change their employer. Similarly to Bils, Shin finds substantially procyclical wages for workers who change employers and much less procyclical wages for job stayers. Solon, Barsky, and Parker (1994) estimate wage cyclicality using data from the PSID for the period 1967–1987. They find that the point estimate of the cyclicality of men's real wages is between -1.35 percent and -1.40 percent. In the sample restricted to workers who did not change employers,

[6] The type of data set, which contains information on the cross section of individuals over time, is called panel data.

the coefficient reduces to -1.24 percent. Consistent with Bils (1985), Solon, Barsky, and Parker find that wages of job stayers are less procyclical than wages of all workers.

The usual measure of wages used in the studies is the average hourly wage, which is constructed by dividing total annual earnings by total annual hours worked. However, if workers are more likely to hold more than one job in expansions, then the constructed average hourly wage may be more procyclical than actual wages within employment relationships. Devereux (2001) conducts a detailed examination of the cyclicality of wages of job stayers and, in contrast to the earlier studies, focuses on the wages of workers who have only one job at a time. Using PSID data from 1970–1992 on men's earnings, Devereux finds that the cyclicality of the average wage of these workers is -0.54 percent. These findings confirm that wages of job stayers are less procyclical than wages of job changers.[7]

Evidence from Matched Firm-Worker Data

Controlling for a firm fixed effect is important if there are changes in the composition of firms over the business cycle with respect to the level of wages they offer. For example, if the firms that hire in economic booms are predominantly high-wage firms and the firms that hire in economic busts are low-wage firms, then the failure to control for the firm's fixed effect may lead to biasing the estimates of the cyclicality away from zero even when wages are acyclical. Researchers often use occupation and industry fixed effects to control for changes in the composition of jobs over the business cycle, which is readily available from worker survey data. Most of the studies employ individual worker survey data that do not allow identification of the firm's fixed effect. To allow identification of a firm fixed effect, the data must contain information on more than one worker from the same firm and on firm identifiers. Only recently have longitudinal firm-worker data for the U.S. economy become available; however, to my knowledge, there are no studies of wage cyclicality using these data yet.

Carneiro, Guimarães, and Portugal (2009) use administrative firm-worker data from Portugal. They estimate a model in levels similar to (17), controlling for an individual worker's qualification, education, age, and a quadratic in time trend. Their findings are very similar to the findings by Bils (1985). In particular, controlling for a worker and a firm fixed effects, they find that the cyclicality of wages of workers who have been with their employer for less than a year is -2.77 percent. The cyclicality of wages of workers who have been with an employer for more than a year, job stayers, is -1.41 percent. Importantly, accounting for both firm and worker fixed effects delivers results

[7] Shin and Solon (2007) find similar evidence in the NLSY data.

very similar to the results from accounting only for worker fixed effect. In particular, the cyclicality of wages of new hires from the regression with only a worker fixed effect is -2.73 percent, while with only a firm fixed effect it is -3.53 percent. The cyclicality of wages of job stayers from the regressions with only a worker fixed effect is -1.50 percent, while from the regression with only a firm fixed effect it is -2.94 percent. Using the same data set, Martins, Solon, and Thomas (2010) investigate the cyclicality of wages of newly hired workers in a subset of occupations into which firms frequently hire new workers. The estimated cyclicality of the wages of newly hired workers in these entry jobs is -1.8 percent. The authors conclude that the wages of new hires in the entry jobs are substantially procyclical.[8]

The results from the studies that allow controlling for firm fixed effects confirm the earlier findings that wages of newly hired workers are more cyclical than wages of existing workers.

Cyclicality of Wages of Job Stayers and Job Changers and Match Quality

Gertler and Trigari (2009) suggest that the difference in the cyclicality of wages of new hires and existing workers can be explained by the differences in the quality (or, alternatively, productivity) of newly formed and ongoing matches. In particular, Gertler and Trigari argue that separately controlling for firm and worker fixed effects cannot account for match quality, which must be controlled for by the interaction term—a worker-job fixed effect. Gertler and Trigari use individual male worker data from the Survey of Income and Program Participation over the period 1990–1996. The data consists of four panels from 1990, 1991, 1992, and 1993, each lasting approximately three years and containing information from interviews conducted every four months. The data allow for identifying if a worker changes employer. Gertler and Trigari estimate a wage equation similar to equation (17) except that, instead of controlling for a worker fixed effect, η_i, they control for the unobserved firm-worker effect, which simultaneously captures two effects: a worker fixed effect that does not change from job to job and a joint worker-firm effect. After the authors control for a worker-firm fixed effect, the coefficient on the interaction term between the unemployment rate and the dummy for new hire becomes small and statistically insignificant. Gertler and Trigari interpret the results as evidence of an omitted variable, a worker-firm specific fixed effect. If a job change is systematically associated with the movement from low to high quality match, then the omitted variable is negatively correlated with the interaction term, biasing the estimates. They conclude that a large part of the

[8] Martins, Solon, and Thomas (2010) conclude that the cyclicality is of the similar magnitude as the cyclical elasticity of employment.

higher procyclicality of wages of new hires is probably due to the comparatively higher quality of these matches. Gertler and Trigari suggest that the existing literature does not provide conclusive evidence that the newly hired workers have more cyclical wages than the existing workers in the same firm.

This finding raises questions, which I discuss at the end of this section, about what heterogeneity in the data should be controlled for when calculating the statistics and how to bring the model to match these statistics.

Dependence of Wages on the Past Labor Market Conditions

Literature surveyed so far finds evidence that the wages of new hires are more sensitive to the aggregate labor market conditions than the wages of workers in ongoing employment relationships. A closer look at workers in ongoing employment relationships shows that their wages depend not only on current labor market conditions, but also on the history of labor market conditions during the entire employment relationship.

Beaudry and DiNardo Regressions

Beaudry and DiNardo (1991) estimate the following equation for individual wages:

$$\ln\left(w_{jt_0t}\right) = X_{jt_0t}\alpha + \gamma_{start}U_{t_0} + \gamma_c U_t + \gamma_{\min}\min\left\{U_\tau\right\}_{\tau=t_0}^{t} + \eta_j + \nu_{jt}, \quad (18)$$

where w_{jt_0t} is an hourly wage of a worker j in year t who was hired in year t_0, X_{jt_0t} is a vector of the individual- and job-specific characteristics, U_τ is the unemployment rate in year τ, η_j is an individual-specific fixed effect, and ν_{jt} is an individual- and time-varying error term. ν_{jt} is assumed to be serially uncorrelated as well as uncorrelated across individuals. The vector of individual- and job-specific characteristics, X_{jt_0t}, includes a quadratic in experience, a quadratic in tenure, years of schooling, and dummies for industry, region, race, union status, marriage, and standard metropolitan statistical area. The equation is estimated using the individual data on men's wages from PSID, 1976–1984, and two cross-sectional samples from the Current Population Survey (CPS).

The main finding of Beaudry and DiNardo (1991) is that when all three measures of the unemployment rate are included, the effect of the minimum unemployment rate is the most significant, both statistically and economically. Thus, whenever the labor market conditions improve, wages increase. In particular, controlling for worker fixed effect in the PSID sample, the coefficient on the minimum unemployment rate is -2.9 percent, the coefficient on the unemployment rate at the start of the job is -0.6 percent and insignificant, and the coefficient on the contemporaneous unemployment rate is -0.7 percent.

If, however, only the contemporaneous unemployment rate is included, then the results are consistent with earlier studies—the coefficient on the contemporaneous unemployment rate is -1.4 percent.

Subsequent studies replicate the findings of Beaudry and DiNardo (1991) for different time periods and using different data sets. McDonald and Worswick (1999) find support in Canadian data. Grant (2003) estimates an equation similar to (18) and adds the maximum unemployment rate experienced by a worker from the start of the job. Grant finds that both the minimum unemployment rate and the contemporaneous unemployment rate have an effect on wages. In particular, in the sample of young men from NLS from 1966–1983, when all three unemployment rates are included, the coefficient on the minimum unemployment rate is -2.29 percent while the coefficient on the contemporaneous unemployment rate is -2.37 percent. This finding leads Grant to conclude that wages depend both on the past and on the contemporaneous labor market conditions.

Devereux and Hart (2007) study the history dependence in wages in British data, the New Earnings Survey Panel, for the period 1976–2001. They estimate a model similar to (18) that also includes the maximum unemployment rate but employ a different estimation procedure from the studies above. They find that both the minimum unemployment rate and the contemporaneous unemployment rate are statistically significant and negative. The authors conclude that the British real wage data exhibit both the history dependence as described in Beaudry and DiNardo (1991) and the dependence on the contemporaneous labor market conditions.

Hagedorn and Manovskii (2009) find that the dependence on the past unemployment rates in model (18) disappears if one controls for the quality of a match. They argue that the quality of a match can be learned from the number of job offers a worker receives throughout the total duration of the job, which can be approximated by the sum of the aggregate vacancy-unemployment ratios experienced by a worker throughout the job. In addition, if a worker switches job-to-job, then the sum of labor market tightnesses experienced during a previous job also helps predict the quality of the current match. Using the NLSY, the authors find that if these controls are included in Beaudry and DiNardo's (1991) equation, (18), the coefficients on the past unemployment rates are insignificant both economically and statistically, while the new controls have a large positive effect.

Evidence from Matched Firm-Worker Data

Baker, Gibbs, and Holmstrom (1994) provide compelling evidence on the history dependence of wages in the study of the wage policy of a large firm over the period 1969–1988. The authors find that there is a substantial cohort effect in wages, where the cohort is defined as the employees who enter the sample in a given year.[9] That is, much of the variation in wages between cohorts comes from the differences in starting wages, which implies that wages depend on the history of the labor market conditions from the start of the job. The authors investigate whether the differences in the starting wage can be driven by observable or unobservable worker characteristics. To check for the possible impact of unobservable characteristics, they examine whether cohorts that entered with lower starting wages are promoted less and exit more. They find no evidence of this and no evidence that composition effect can fully account for either the differences in starting wages or the persistent effect of external labor market conditions from the start of the job on wages.

Using a large matched employer-employee data set from Northern Italy, Macis (2006) provides a detailed empirical investigation of the dependence of wages on the unemployment rates from the start of the job, controlling for both firm and worker fixed effects. Using a model similar to (18), Macis finds that wages are correlated with both the best and the worst labor market conditions from the start of the job, as well as with the contemporaneous unemployment rate.

Cyclicality of a Measure of Wages that Takes into Account Future Wages

The studies reviewed above estimate the cyclicality of the *current* wage. These studies find that wages of newly hired workers are more procyclical than wages of workers in ongoing employment relationships, and wages depend on the history of labor market conditions from the start of the job. As discussed earlier, what is relevant for job creation is the expected present discounted value of wages paid in a newly formed match. Nevertheless, from the evidence presented so far we can form some intuition about the cyclical behavior of the measure of wages that takes into account both the initial wage and the expected value of future wages.

Consider a firm that decides whether to hire a worker in the current period or to hire in the following period. In addition, suppose that in the current period unemployment is high but is expected to return back to its lower level in the following period. Since wages of newly hired workers are procyclical, the hiring wage in the employment relationship that starts in the current period is low. Because of the history dependence of wages, the future wages in this relationship are also expected to be lower than the wages in the matches formed in the future periods. Thus, by hiring now a firm locks in a worker to a stream

[9] In the study, the authors cannot identify whether the entrants are the new hires at the firm or are internally promoted. They argue that it is plausible that both categories of workers are treated in the same way by the firm. Their comparison of wage patterns of these workers with industry wages supports this view.

of wages that is lower as compared to a stream of wages to be paid to a worker hired the following period. Consequently, the wage costs associated with hiring in the current period are comparatively lower because the initial wage is low and because of the future wage savings. Similarly, if in the current period the unemployment rate is low and is expected to increase in the following period, the total wage costs associated with hiring in the current period are comparatively higher than the total wage costs associated with hiring in the following period. This argument, developed in Kudlyak (2009), suggests that the relevant measure of wages that a firm takes into account at the time of hiring is low when unemployment is high and high when unemployment is low, which is the opposite of being rigid. To gauge the quantitative importance, we need empirical estimates of this cyclical volatility.

Using the free entry condition for firms, Haefke, Sonntag, and van Rens (2009) argue for the importance of the elasticity of the expected present discounted value of wages with respect to the expected present discounted value of productivity in newly formed matches, which they refer to as permanent values of wages and productivity, respectively. They do not estimate the elasticity directly but aim at providing the empirical bounds for this statistic. Using simulations of the standard model, Haefke, Sonntag, and van Rens conclude that "the elasticity of the current period wage of newly hired workers with respect to current period productivity . . . constitutes a good proxy for the elasticity of the permanent wage with respect to permanent productivity for the case of instantaneously rebargained wages." Using the simulations of a model, similar to the modified model presented in Section 1 above, they argue that "the elasticity of the current period wage of newly hired workers with respect to current period productivity . . . can be seen as a lower bound for [the elasticity of the permanent wage with respect to permanent productivity] in the case of wage rigidity on the job." Haefke, Sonntag, and van Rens (2009) proceed to estimate the elasticity of wages of newly hired workers with respect to productivity using a large data set on wages of newly hired workers from nonemployment from the CPS. The estimated model for wages is similar to the models presented above except that, instead of using the series of unemployment, they use the series of labor productivity as a cyclical indicator. They find that the elasticity of wages of newly hired workers from nonemployment, 0.8, is substantially larger than the elasticity of wages of all workers, 0.2.[10]

Kudlyak (2009) provides an estimate of the cyclical behavior of the measure of wages that takes into account the initial wage and the expected present value of future wages to be paid in a newly formed match. The firm's hiring decision can be thought of as a decision to hire in the current period versus

[10] Haefke, Sonntag, and van Rens (2009) document that the elasticity of wages of job-to-job movers is similar to the elasticity of wages of newly hired workers from nonemployment or even larger.

waiting one more period and hiring then. In equilibrium, the marginal productivity of an additional worker equals the user cost of labor, which is the difference between the expected present discounted values of the costs associated with creating a match with a worker in the current period and the costs associated with creating a match the following period. In a model with search and matching, these costs consist of expenses on hiring a worker, i.e., costs associated with vacancy posting, and wage payments to a worker.

Using individual wage data from the NLSY, Kudlyak (2009) estimates the cyclicality of the wage component of the user cost of labor, which equals the wage at the time of hiring plus the expected present discounted value of the differences from the next period onwards between the wages paid to the worker hired in the current period and the worker hired the following period.[11] The estimated cyclicality of the wage component of the user cost of labor is -4.5 percent as compared to the cyclicality of wages of newly hired workers of -3 percent. The greater cyclicality obtains because at the time of hiring, a firm to some degree locks in a worker to a stream of wages that depends on the economic conditions from the start of the job. Thus, the rigidity of wages within employment relationships actually amplifies the fluctuations of the expected present discounted value of wages to be paid to a newly hired worker as compared to the fluctuations in the initial wage in newly formed matches.

Discussion

To gauge whether the wage data exhibit enough rigidity to amplify fluctuations in the standard search and matching model, the empirical estimates from wage data must be contrasted with the statistics obtained from the model. This task is conducted in Pissarides (2009), Haefke, Sonntag, and van Rens (2009), and Kudlyak (2009).

Pissarides (2009) compares the elasticity of wages with respect to productivity obtained from the standard search and matching model using common parameter values to the elasticity of wages of newly hired workers with respect to productivity in the data. He finds that the elasticity of wages of new hires with respect to productivity in the data is close to 1. This is consistent with the elasticity of wages generated by the standard search and matching model with Nash bargaining. He concludes that any solution to the unemployment volatility puzzle should be able to generate this near-proportionality of wages of new hires and productivity. Thus, a model with more rigid wages will not be able to match the data. The same conclusion is reached by Haefke, Sonntag, and van Rens (2009), who compare their estimates of the elasticity of wages

[11] See Kudlyak (2007) for more details on the estimation.

of newly hired workers with respect to productivity to the statistics from the standard search and matching model.

Kudlyak (2009) calibrates the elasticity of the wage component of the user cost of labor in the model to the empirical estimate and examines how much volatility of vacancies and unemployment the calibrated model can generate. She concludes that the data lack required rigidity to amplify the fluctuations of vacancies and unemployment in the model.

Statistics Conditional on Match Quality

Gertler and Trigari (2009) find that, conditional on match quality, there is no difference between the cyclicality of wages of new hires and existing workers. They argue that their finding implies that assuming the same cyclicality for new hires' and existing workers' wages within each firm in the standard search and matching model is consistent with the existing micropanel data evidence on new hires' wages once the empirical evidence controls for match quality, i.e., match productivity. Their finding that, conditional on match quality, there is no difference between the cyclicality of wages of new hires and existing workers is consistent with the evidence that the wages that firms pay to newly hired workers are (unconditionally) more procyclical than the wages of workers in ongoing matches. Note, however, that if the conditional statistics are used for calibrating the model, i.e., if the wage statistics from the model are compared to the conditional wage statistics in the data, then the driving force of the model—productivity—as well as other statistics in the model should also be conditioned accordingly.

Gertler and Trigari's evidence suggests a possible source of the difference between the cyclicality of wages of new hires and wages of the existing workers—the difference between the quality of a newly formed match and of an existing match. It implies that there is economically significant cyclical heterogeneity in match quality between newly formed and ongoing matches. In contrast, in the standard search and matching model, newly formed and ongoing matches are homogeneous, i.e., they are equally productive in every period. Thus, for the model to generate cyclical volatilities, the finding calls for a modification of the model to incorporate the cyclical heterogeneity.

3. CONCLUSION

What matters for the hiring decision of a firm over the business cycle is the cyclicality of the expected value of wages paid in newly formed matches. Most of the existing studies are concerned with the cyclicality of the current wage. The evidence on the cyclicality of the expected present value of future wages to be paid in a newly formed match is scarce.

The data provide evidence of the difference between the cyclicality of wages of newly hired workers and of wages of workers in ongoing matches.

In particular, the studies document that a one percentage point increase of the unemployment rate is associated with approximately a 3 percent decrease in wages of newly hired workers. Wages of workers in ongoing matches are less responsive to the contemporaneous labor market conditions and depend on the history of the labor market conditions from the start of the job, i.e., they are more rigid as compared with the wages of newly hired workers. This wage rigidity within employment relationships may, in fact, make the expected present value of wages to be paid in newly formed matches more cyclically volatile than the wage of new hires.

Haefke, Sonntag, and van Rens (2009), using simulations from the model, argue that the wage measure that takes into account future wages in a match is likely more volatile than wages of new hires. Kudlyak (2009) provides an estimate of the cyclicality of the user cost of labor, which takes into account hiring wage and the expected future wages to be paid in the employment relationship. She finds that a one percentage point increase in the unemployment rate is associated with a 4.5 percent decrease in the expected difference between the present value of wages to be paid in a match created in the current period and in a match created in the following period.

The evidence suggests that the measure of wages relevant for job creation is rather procyclical. In fact, using the existing empirical evidence and also providing new estimates, recent studies find that, quantitatively, the data may not exhibit the required rigidity necessary to generate the empirical volatility of unemployment in the standard search and matching model.

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Changes in Monetary Policy and the Variation in Interest Rate Changes Across Credit Markets

Devin Reilly and Pierre-Daniel G. Sarte

The conduct of monetary policy is most often interpreted in terms of the federal funds target rate set by the Federal Open Market Committee (FOMC), at least until recently when this rate effectively reached its zero bound and additional actions were then implemented. The federal funds rate is the interest rate at which private depository institutions, typically banks, lend balances held with the Federal Reserve to other depository institutions overnight. By targeting a particular value for that rate, the Federal Reserve seeks to adjust the liquidity provided to the banking system through daily operations. Because the federal funds rate applies to overnight transactions between financial institutions, it represents a relatively risk-free rate. As such, it serves to anchor numerous other interest rates that reflect a wide array of credit transactions throughout the U.S. economy, such as deposits, home loans, and corporate loans.

Because the federal funds rate anchors interest rates in many different types of credit transactions, monetary policy actions that move the funds rate in a given direction are expected to move other interest rates in the same general direction. However, the extent to which changes in the federal funds rate affect conditions in different credit markets may vary significantly from market to market. For example, changes in the federal funds rate may be closely linked to changes in the three-month Treasury bill rate, but potentially less so to changes in home loan rates. In that sense, changes in monetary policy,

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as reflected by broad liquidity adjustments through the federal funds market, will be more effective in influencing credit conditions in some markets than others. Thus, this article attempts to assess empirically the extent to which interest rate changes in various credit markets reflect changes in monetary policy. It also explores whether these relationships have changed over time.

As a first step, we construct a panel of 86 time series spanning a diverse set of monthly interest rate changes, including Treasury bill rates, corporate interest rates, repurchase agreement rates, and mortgage rates, among others. The panel of interest rate changes covers the period July 1991–December 2009. The empirical framework then uses principal component analysis to characterize co-movement across these interest rate changes. The basic intuition underlying the exercise is as follows: If changes in monetary policy tend to move a broad array of interest rates in the same general direction, then changes in these interest rates will share some degree of co-movement.

Having characterized the common variation in interest rates using principal components, we ask two questions. First, looking across all interest rate changes, which series tend to be mostly driven by common changes in interest rates rather than idiosyncratic considerations? In particular, idiosyncratic changes in a given interest rate series are orthogonal to the principal components and, therefore, unlikely to reflect a common element such as a change in monetary policy. Therefore, one expects that monetary policy will have only a limited effect on interest rates in which changes are mostly idiosyncratic. Second, recognizing that the common variation across interest rate changes may reflect a broad set of aggregate factors, how closely is the common change component of each interest rate series (which may play a more or less important role in the characterization of different interest rates) related to changes in monetary policy? Furthermore, has this relationship changed over time?

Our results indicate that most of the variation across our sample can be explained by a small number of common components. For most credit markets, including mortgage, repurchase agreement, Treasury, and London Interbank Offered Rate (LIBOR) rates, four components explain approximately 70 percent or more of the variation in interest rate changes. One notable exception is the auto loan market, in which interest rate variation is almost entirely idiosyncratic. For most of the series in our sample, the common variation in interest rate changes is relatively highly correlated with the federal funds rate. This suggests that common movements in interest rates reflect, to a large extent, changes in monetary policy, as defined by the federal funds rate, rather than other aggregate disturbances. That said, there nevertheless remains a moderate number of rates for which the common components, while explaining a significant portion of their variability, are not highly correlated with the federal funds rate. We interpret this finding in mainly two ways. First, these rates, which include corporate bond and mortgage rates, are driven to a greater degree by aggregate factors that may be somewhat disconnected from monetary policy. Second, these rates, to the extent that they include longer term rates, reflect monetary policy more indirectly through changes in expected future short rates. For example, changes in beliefs regarding future productivity will likely affect the perceived path of future federal funds rates.

The rest of this article is organized as follows. In Section 1, we review the relevant literature. Section 2 outlines the principal component methodology and calculations used in our analysis. Section 3 describes the data set used in the empirical work. Section 4 presents our findings, while Section 5 offers concluding remarks.

1. LITERATURE REVIEW

Several recent papers have utilized principal component analysis or similar techniques to explore the behavior of various interest rates and macroeconomic variables over time. Diebold, Rudebusch, and Aruoba (2006) and Bianchi, Mumtaz, and Surico (2009), among others, use a latent factor model to explore the interaction between yield curves and several macroeconomic variables, including a monetary policy instrument. The approach used is similar to using principal components to obtain the factors; however, it differs in that principal component analysis requires factors to be orthogonal to each other, but remains agnostic about the form of the factor loadings. The models used in these and other papers restrict the factor loadings by extending an approach for modeling yield curves from Nelson and Siegel (1987). These papers have also restricted their attention to government bond yields.

Perhaps more closely related to our paper is Knez, Litterman, and Scheinkman (1994). This article investigates the behavior of money market instruments utilizing a factor model that is less restrictive on the loadings than the previous papers discussed. The authors find that much of the total variation in their data set can be explained by three or four factors, and that each factor can be interpreted as a parameter that characterizes systematic movements in the yield curve. It differs from our analysis in that the data set used is much narrower, including only Treasury bills, commercial paper, certificates of deposit, Eurodollar deposits, and bankers' acceptances, all with maturities of less than one year. Additionally, they examine the returns of these securities, whereas we analyze the changes in interest rates across a variety of credit markets.

Finally, Gürkaynak, Sack, and Swanson (2005) and Reinhart and Sack (2005) examine the immediate impact of a variety of forms of FOMC communication on several financial variables, including interest rates, equity prices, and others. They use principal components to extract common components from a set of changes in these variables around FOMC statements, testimonies, and other releases. They find that a small number of factors appears to explain a significant amount of the variation in response to all types of FOMC communication. Our analysis does not limit itself to changes in interest rates around FOMC communication, and explores a broader array of rates than these two articles.

2. PRINCIPAL COMPONENTS

Consider a panel of (demeaned) observations on interest rate changes across N credit markets over T time periods, which we summarize in an $N \times T$ matrix, X. Let X_t denote a column of X (i.e., a set of observations on all interest rate changes at date t). As explained in Malysheva and Sarte (2009), the nature of the principal component problem is to ask how much independence there really is in the set of N variables. To this end, the principal component problem transforms the Xs into a new set of variables that will be pairwise uncorrelated and of which the first will have the maximum possible variance, the second the maximum possible variance among those uncorrelated with the first, and so on.

We denote the jth principal component of X by f_j, where

$$f_j = \lambda_j' X, \tag{1}$$

and λ_j' and f_j are $1 \times N$ and $1 \times T$ vectors, respectively. In other words, different principal components of X simply reflect different linear combinations of interest rate changes across sectors. Moreover, the sum of squares of a given principal component, f_j, is

$$f_j f_j' = \lambda_j' \Sigma_{XX} \lambda_j, \tag{2}$$

where $\Sigma_{XX} = XX'$ represents the variance-covariance matrix (when divided by T) of interest rate changes in the data set.

Let $\Lambda_k = (\lambda_1, ..., \lambda_k)$ denote an $N \times k$ matrix of weights used to construct the first k principal components of X, $f_1, .., f_k$, which we arrange in the $k \times T$ matrix $F_k = (f_1', ..., f_k')'$. Thus, $F_k = \Lambda_k' X$ and the principal component problem is defined as choosing sets of weights, Λ_k, that solve

$$\max_{\Lambda_k} \ \Lambda_k' \Sigma_{XX} \Lambda_k \text{ subject to } \Lambda_k' \Lambda_k = I_k. \tag{3}$$

The solution to the above problem has the property that each set of weights, λ_j, solves[1]

$$\Sigma_{XX} \lambda_j = \mu_j \lambda_j, \tag{4}$$

where $\lambda_j' \lambda_j = 1 \, \forall j$. Put another way, the sets of weights that define the different principal components of X in equation (1) are eigenvectors, λ_j, of the the variance-covariance matrix of interest rate changes, Σ_{XX}, with corresponding eigenvalues given by μ_j. In addition, because the variance-covariance matrix

[1] See the Appendix in Malysheva and Sarte (2009).

of X is symmetric, these eigenvectors are orthogonal to each other, $\lambda_j'\lambda_i = 0$ $\forall i \neq j$.

Combining equations (2) and (4), note that

$$f_j f_j' = \lambda_j' \mu_j \lambda_j = \mu_j. \tag{5}$$

Therefore, the eigenvalue μ_j is the sum of squares of the principal component f_j in (2). Then, given that principal components are ranked by the extent of their variance, the first such component, f_1, is obtained using the weights, λ_1', associated with the largest eigenvalue of Σ_{XX}. The second principal component is obtained using the weights corresponding to the second largest eigenvalue of Σ_{XX}, and so on.

Proceeding in this way for each of the N principal components of X using the weights given by (4), observe that

$$\Lambda_N' \Sigma_{XX} \Lambda_N = \begin{bmatrix} \mu_1 & 0 & ... & 0 \\ 0 & \mu_2 & ... & 0 \\ 0 & 0 & ... & \mu_N \end{bmatrix}. \tag{6}$$

If the rank of Σ_{XX} were $k < N$, there would be $N - k$ zero eigenvalues and the variation in interest rate changes would be completely captured by k independent variables. In fact, even if Σ_{XX} has full rank, some of its eigenvalues may still be close to zero so that a small number of (or the first few) principal components may account for a substantial proportion of the variance of interest rate changes.

The Appendix at the end of the article shows that the principal component problem defined in (3) can be derived as the solution to the least square problem

$$\min_{\{f_1,\dots,f_k\}_{t=1}^{T},\Lambda_k} T^{-1} \sum_{t=1}^{T} e_t' e_t \text{ subject to } \Lambda_k'\Lambda_k = I_k, \tag{7}$$

where

$$X_t = \Lambda_k F_{k,t} + e_t. \tag{8}$$

Hence, it follows that

$$\Sigma_{XX} = \Lambda_k \Sigma_{FF} \Lambda_k' + \Sigma_{ee}, \tag{9}$$

where $\Sigma_{FF} = F_k F_k'$, in which case we can think of the principal components as capturing some portion $\Lambda_k \Sigma_{FF} \Lambda_k'$ of the variation in interest rate changes, Σ_{XX}.

Given the decomposition expressed in (8), each interest rate change in the data set can be written as

$$\Delta r_t^i = \Lambda_k^i F_{k,t} + e_t^i, \tag{10}$$

where Λ_k^i is the ith row of Λ_k. In that sense, $\Lambda_k^i F_{k,t}$ captures the importance of the principal components in driving each individual series. The objective of the article then is to address two key aspects of interest rate changes.

First, having computed a set of principal components, $F_{k,t}$, that account for most of the fraction of the variation in the Xs, we wish to assess the extent to which a given series of interest rate changes, Δr_t^i, is driven by these components rather than its own disturbance term, e_t^i. The important consideration here is that the principal components, $F_{k,t}$, in (10) are common to all interest rate changes (i.e., they do not depend on i) and, therefore, will be directly responsible for co-movement across interest rate changes. In contrast, even if there remains some covariation across the shocks, e_t^i, this covariation will, by construction, play a larger role in explaining idiosyncratic variations in interest rate changes. In that sense, changes in monetary policy will more likely be reflected in the co-movement term $\Lambda_k^i F_{k,t}$ in equation (10) rather than e_t^i.

Formally, we compute how much of the variance of Δr_t^i, denoted $\sigma^2_{\Delta r_{it}}$, is explained by the variance of $\Lambda_k^i F_{k,t}$,

$$R_i^2(F) = \frac{\Lambda_k^i \Sigma_{FF} \Lambda_k^{i\prime}}{\sigma^2_{\Delta r_{it}}}. \tag{11}$$

The series of interest rate changes with $R_i^2(F)$ statistics close to 1 are driven almost entirely by forces that determine mainly the covariation across interest rate changes. In contrast, series of interest rate changes with $R_i^2(F)$ statistics close to zero generally reflect considerations that are likely more idiosyncratic to each series.

Suppose that we were interested in a subgroup of M series—say all mortgage interest rates or all repurchase agreement rates. We can compute an analogous R^2 statistic for that credit market segment by using a $1 \times N$ weight vector, $\mathbf{w}$, that associates positive weights to the series of interest and zeros elsewhere. The implied $R_M^2(F)$ statistic is then given by

$$R_M^2(F) = \frac{\mathbf{w}\Lambda \Sigma_{FF} \Lambda' \mathbf{w}'}{\mathbf{w}\Sigma_{XX}\mathbf{w}'}. \tag{12}$$

As before, $R_M^2(F)$ statistics close to 1 indicate a subgroup of credit markets (defined by the weights, $\mathbf{w}$) that are mostly affected by common forces across interest rates, $\mathbf{w}\Lambda_k F_{k,t}$, rather than conditions specific to that subgroup, $\mathbf{w}e_t$.

Second, because the covariation across interest rate changes reflects not only changes in monetary policy but also other aggregate considerations (including those potentially driven by systemic issues), the next step is to relate changes in each interest rate series captured by principal components to changes in monetary policy. Hence, in each credit market, i, we compute the correlation between $\Lambda_k^i F_{k,t}$ and changes in the effective federal funds rate, Δr_t^{fed},

$$\rho_i = \mathrm{corr}(\Lambda_k^i F_{k,t}, \Delta r_t^{fed}).$$

Evidently, Δr_t^{fed} may not capture all of the relevant aspects of changes in monetary policy and serves here only as an approximate guide. For instance, going forward, we may be more interested in the relationship between $\Lambda_k^i F_{k,t}$ and the interest on reserves. More generally, to the extent that other measurable aspects of changes in monetary policy matter, say represented in a vector Z_t, one could instead compute the projection,

$$\Lambda_k^i F_{k,t} = Z_t \beta + u_t^i,$$

and its associated $R_i^2(Z)$ statistic.

3. THE DATA

Our analysis focuses on a data set that includes 86 time series on interest rate changes, all seasonally adjusted and expressed at an annual rate. These include a wide array of rates with monthly observations spanning back to July 1991. A full list of rates and associated descriptive statistics can be found in the Appendix (Table 5). The data come primarily from Haver Analytics and Bloomberg. While we analyze these rates individually, for ease of presentation we also place them into eight broad categories and investigate the average behavior in each of these credit markets.

The first group includes LIBOR rates based on the U.S. dollar, with maturities ranging from one month to one year. These are reference rates based on the interest rates at which banks are able to borrow unsecured funds from other banks in the London interbank market. We refer to the second group in our data set as the deposit group, which contains averages of dealer offering rates on certificates of deposit with maturities from one to nine months, as well as bid and effective rates on Eurodollar deposits for maturities of overnight to one year. Our third group includes a variety of Treasury bill, note, and bond rates. There are two secondary market rates (three- and six-month), which are the average rates on Treasury bills traded in the secondary market. We also include auction highs on three- and six-month Treasury bills. However, the majority of rates in this group are yields on nominal Treasury securities with maturities ranging from three months to 30 years. These are interpolated by the U.S. Treasury from the daily yield curve for noninflation indexed securities, based on closing market bid yields on actively traded Treasury securities.

Our panel also contains a variety of corporate borrowing rates. We include one-month and three-month rates for nonfinancial and financial commercial paper in this group. These rates are calculated by the Federal Reserve Board using commercial paper trade data from the Depository Trust and Clearing Corporation. Also included are Aaa and Baa Moody's corporate bond yields, which are based on outstanding corporate bonds with remaining maturities of at least 20 years. Finally, Citigroup Global Markets provides corporate bond yields that cover a variety of industries and ratings.

We include two smaller groups, one of which contains three rates for long-term government (state and local) and agency bonds. The other relatively small group in our panel includes two series of interest rate changes for new and used car loans. These are simple unweighted averages of rates commonly charged by commercial banks on auto loans.

The final two groups we utilize are mortgage rates and repurchase agreement rates. The former spans a variety of mortgage rates, including new homes, existing homes, adjustable rate loans, and fixed rate loans. The repurchase agreement group (which also includes reverse repurchase rates) is based on transactions that involve Treasury, mortgage-backed, or agency securities, with maturities ranging from one day to three months.

4. EMPIRICAL FINDINGS

Given the computation of principal components described in Section 1, the next section assesses the extent to which a small number of principal components, out of potentially 86, captures the variation in interest rates across different credit markets. We then gauge the contribution of common changes to individual interest rate variations, as captured by the $R_i^2(F)$ statistic described above. In other words, in each credit market, we assess how much of the variation in its interest rate, Δr_t^i, is explained by its component related to common interest rate movements, $\Lambda_k^i F_{k,t}$. The next subsection then relates the common component of individual interest rate changes, $\Lambda_k^i F_{k,t}$, to changes in the federal funds rate, Δr_t^{fed}, by examining their correlation, $\mathrm{corr}(\Lambda_k^i F_{k,t}, \Delta r_t^{fed})$. Finally, in the last subsection, we examine the robustness of our findings over different sample periods.

Accounting for Interest Rate Variations with a Small Number of Factors

This subsection examines the degree to which a small number of factors potentially captures most of the variation in interest rates across credit markets. We carry out this assessment in mainly two ways. First, we ask how much of the variation in average interest rate changes, $N^{-1} \sum_{i=1}^{N} \Delta r_t^i$, is explained by the first few principal components. Second, following Johnston (1984), we ask how much of the sum of individual variations in the Xs is explained by these components. The total individual variation in interest rate changes is given by

$$\sum_{t=1}^{T} (\Delta r_t^1)^2 + \sum_{t=1}^{T} (\Delta r_t^2)^2 + ... + \sum_{t=1}^{T} (\Delta r_t^N)^2 = tr(\Sigma_{XX}). \tag{13}$$

From equation (6), observe that

$$\begin{aligned}\sum_{i=1}^{N}\mu_j &= tr(\Lambda_N' \Sigma_{xx}\Lambda_N) \\ &= tr(\Sigma_{XX}\Lambda_N\Lambda_N') \\ &= tr(\Sigma_{XX}). \end{aligned} \qquad (14)$$

In other words, the sum of the eigenvalues of the covariance matrix of interest rate changes, Σ_{XX}, is precisely the sum of individual variations in these changes. It follows that

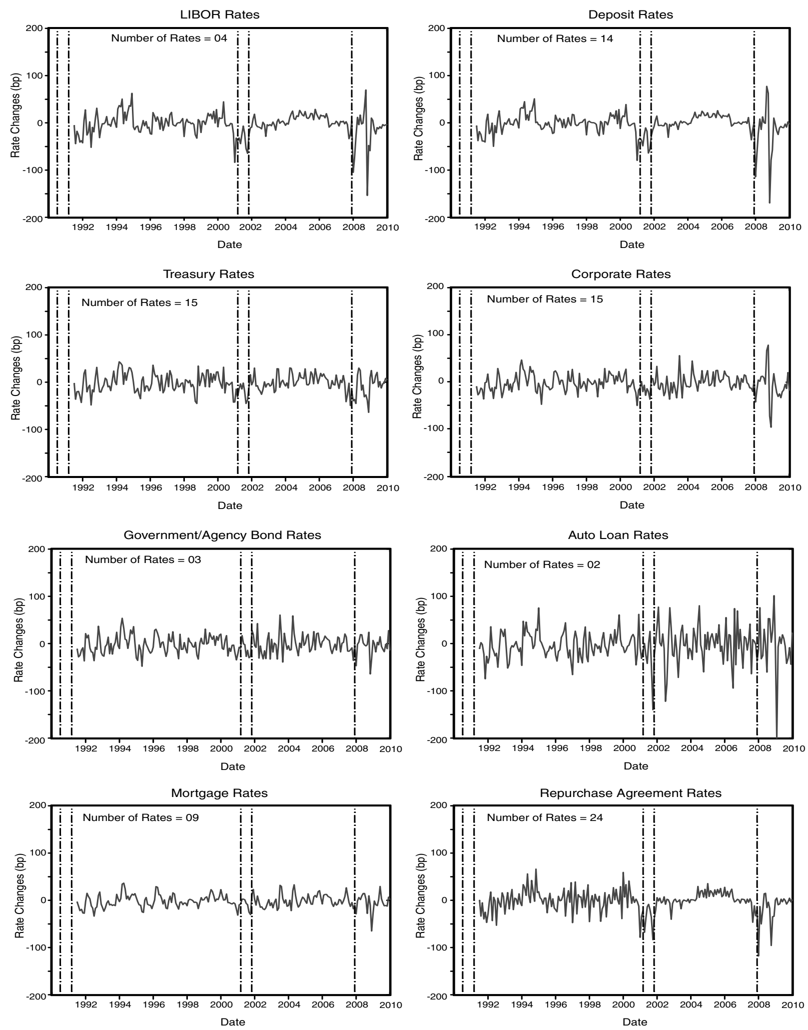
$$\frac{\mu_1}{\sum_{i=1}^{N}\mu_j}, \frac{\mu_2}{\sum_{i=1}^{N}\mu_j}, ..., \frac{\mu_N}{\sum_{i=1}^{N}\mu_j} \qquad (15)$$

represent the proportionate contributions of each principal component to the total individual variation in interest rate changes. In addition, since principal components are orthogonal, these proportionate contributions add up to 1.

The analysis reveals that the first four principal components (i.e., $k = 4$) of the panel of interest rate changes constructed for this article explain 99 percent of the variation in average interest rate changes, $N^{-1}\sum_{i=1}^{N}\Delta r_t^i$, and 78 percent of the their total individual variation, $\frac{\sum_{k=1}^{4}\mu_k}{\sum_{i=1}^{N}\mu_j} = 0.78$. In other words, a small number of components effectively accounts for the variation in the data set. The findings discussed in the remainder of the article are based on these first four principal components. However, our conclusions regarding the effects of changes in monetary policy in different credit markets, in particular the qualitative ranking of credit markets most influenced by changes in the federal funds rate, are robust to considering either fewer than four or up to eight principal components.

As discussed in the prior section, we summarize the behavior of our interest rate series into eight main categories. Figure 1 depicts average changes in these eight broad credit markets over time. Recession peaks and troughs are indicated in the figures by vertical dashed lines. The average changes in rates differ in both persistence and volatility across the eight groups. At two extremes, changes in mortgage rates appear to be relatively stable relative to other rates, whereas auto loan rates are considerably more volatile than any other group. Table 1A provides basic summary statistics for each category of credit markets, as well as for the effective federal funds rate. Consistent with Figure 1, Table 1A indicates that auto loan rates are by far the most volatile rates while mortgage rates are least volatile. In addition, many of these interest rate changes, including auto loan, deposit, and mortgage rates, present evidence of kurtosis. That is, much of the variance in these interest rate changes stems from infrequent extreme observations as opposed to relatively common deviations. Some of the series also show evidence of skewness. For

Figure 1 Average Interest Rates in Different Credit Market Segments



example, deposit, auto loan, and LIBOR rates are all left skewed, indicating the presence of large negative changes in the time series.

Table 1B presents analogous summary statistics for individual Treasury bill rates of different maturities. Interestingly, the standard deviations of the rates increase for maturities of three months to three years, and then decrease at higher maturities. In addition, shorter-term rates, namely three months and six months, are left skewed and thus have historically experienced large

Table 1 Changes in Rates by Category

Table 1A: Changes in Rates, by Credit Market

Series	Mean	Std. Dev.	Skewness	Kurtosis	Min.	Max.
Federal Funds	−2.60	19.87	−1.27	6.49	−96	53
LIBOR	−2.68	25.67	−1.62	12.98	−219	94
Deposit	−2.59	27.38	−1.90	17.87	−285	140
Treasury	−2.35	22.71	−0.33	4.08	−111	65
Corporate	−2.24	26.10	0.34	11.54	−179	227
Government/Agency	−2.14	22.72	0.05	3.99	−86	77
Auto	−3.73	48.41	−1.37	15.58	−392	172
Mortgage	−1.94	19.45	0.82	15.03	−110	200
Repurchase Agreements	−2.56	30.03	−0.98	9.13	−225	168

Table 1B: Changes in Treasury Rates, by Maturity

Series	Mean	Std. Dev.	Skewness	Kurtosis	Min.	Max.
Three-Month Bill	−2.55	21.96	−1.06	4.99	−89	49
Six-Month Bill	−2.63	22.16	−0.69	4.11	−77	54
One-Year Bill	−2.70	23.41	−0.43	3.66	−79	60
Two-Year Note	−2.70	36.06	−0.03	2.86	−69	63
Three-Year Note	−2.65	26.85	0.11	2.76	−69	65
Five-Year Note	−2.42	26.06	0.13	2.85	−77	60
Seven-Year Note	−2.22	24.63	0.13	3.28	−93	61
10-Year Note	−2.04	23.43	−0.08	4.35	−111	65
20-Year Bond	−1.74	20.75	−0.18	5.37	−109	58
30-Year Bond	−1.74	19.86	−0.36	6.00	−110	51

Notes: Basis points, monthly at annual rate.

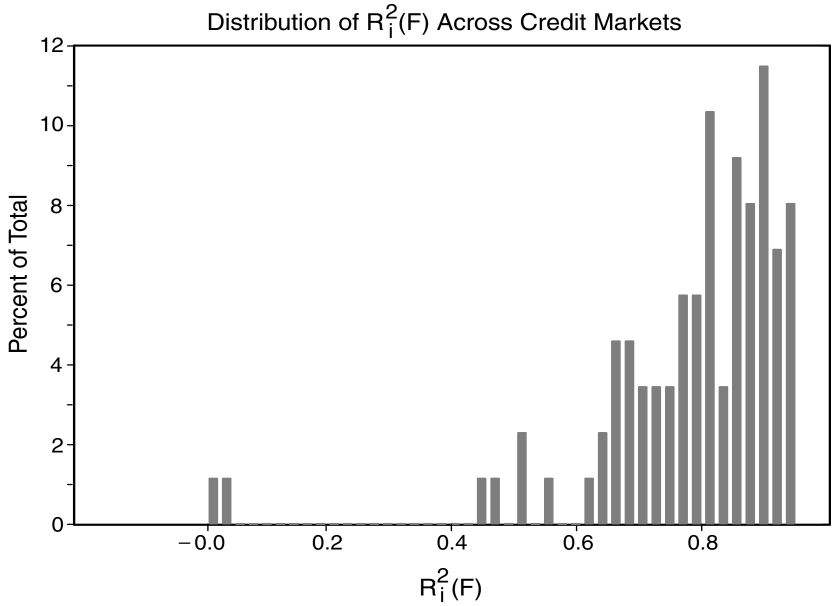
negative changes. Finally, changes in Treasury rates with maturities less than one year and more than 10 years also have relatively large kurtosis statistics.

Contribution of Common Changes to Individual Interest Rate Variations

Figure 2 shows a histogram of the $R_i^2(F)$ statistic discussed in Section 1. This statistic captures the extent to which common movements across interest rates, as summarized by $\Lambda_k^i F_{k,t}$ for each individual interest rate, drive changes in these individual rates. Two main observations stand out. First, changes in interest rates across credit markets tend to reflect factors common to all interest rate changes. In particular, the median $R_i^2(F)$ statistic is 0.814. Second, this first observation notwithstanding, the data also include interest rates in which variations appear almost exclusively driven by idiosyncratic considerations rather than common changes. This is true, for example, of auto loan rates.

Table 2A presents the $R_M^2(F)$ statistics for the eight broad categories of credit markets described earlier. These range from 0.03 for auto loan rates

Figure 2 Importance of Common Changes in Individual Interest Rates



to 0.92 for LIBOR rates.[2] In other words, changes in auto loan rates are explained almost exclusively by idiosyncratic considerations. Put another way, factors that explain co-movement across interest rate changes, one of which is expected to be monetary policy, appear to have little influence over interest rate variations in the auto loan credit market. At the other extreme, changes in LIBOR and deposit rates are almost exclusively driven by forces responsible for the co-movement across interest rates. Somewhere between these two extremes, observe that the common components explain about 68 percent of the variation in government and agency bond rates and mortgage rates.

Table 2B presents the same $R_i^2(F)$ statistics for Treasury bill rates of different maturities. As indicated in the table, the principal components play a large role in explaining the variation in these rates across all maturities. In this case, the $R_i^2(F)$ statistics range from 0.71 to 0.91. Around 78 percent of the variation in 30-year Treasury bill rates is explained by forces common to all interest rates. Interestingly, the common component of the three-month

[2] A listing of all $R_i^2(F)$ statistics can be found in the Appendix (Table 6).

Table 2 Importance of Principal Components in Different Interest Rate Categories

Table 2A: Average $R^2_M(F)$ by Credit Market Segment

Series	**Average $R^2_i(F)$**
Auto	0.030
Mortgage	0.682
Government/Agency	0.685
Repurchase Agreements	0.754
Treasury	0.835
Corporate	0.839
Deposit	0.860
LIBOR	0.917

Table 2B: Average $R^2_i(F)$ for Treasury Securities

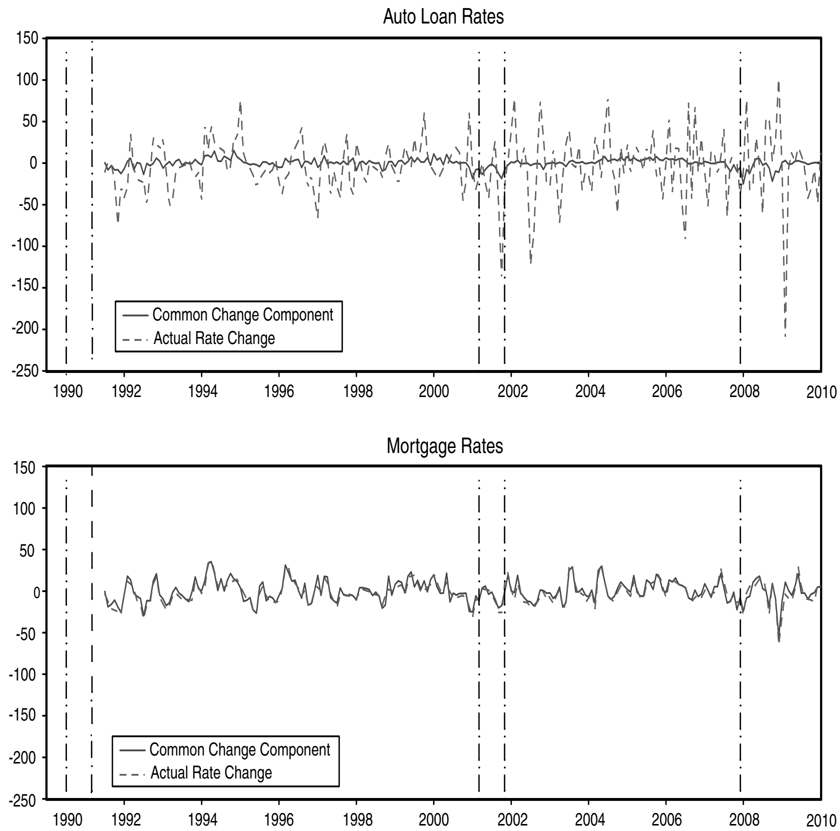
Series	**Average $R^2_i(F)$**
Three-Month Bill	0.710
Six-Month Bill	0.846
One-Year Bill	0.866
Two-Year Note	0.856
Three-Year Note	0.873
Five-Year Note	0.902
Seven-Year Note	0.913
10-Year Note	0.907
20-Year Bond	0.844
30-Year Bond	0.781

Notes: Monthly rates.

Treasury bill rates explains a lower fraction of the variation in that rate than does the corresponding common component in the 30-year rate. However, since common sources of movement in Treasury bill rates reflect changes not only in monetary policy but also in other aggregate factors, one cannot conclude from Table 2B that changes in the federal funds rate exert a greater influence on the 30-year rate than the three-month rate. For the same reason, it does not follow from Table 2B that changes in monetary policy broadly affect Treasury bill rates to the same degree across all maturities.

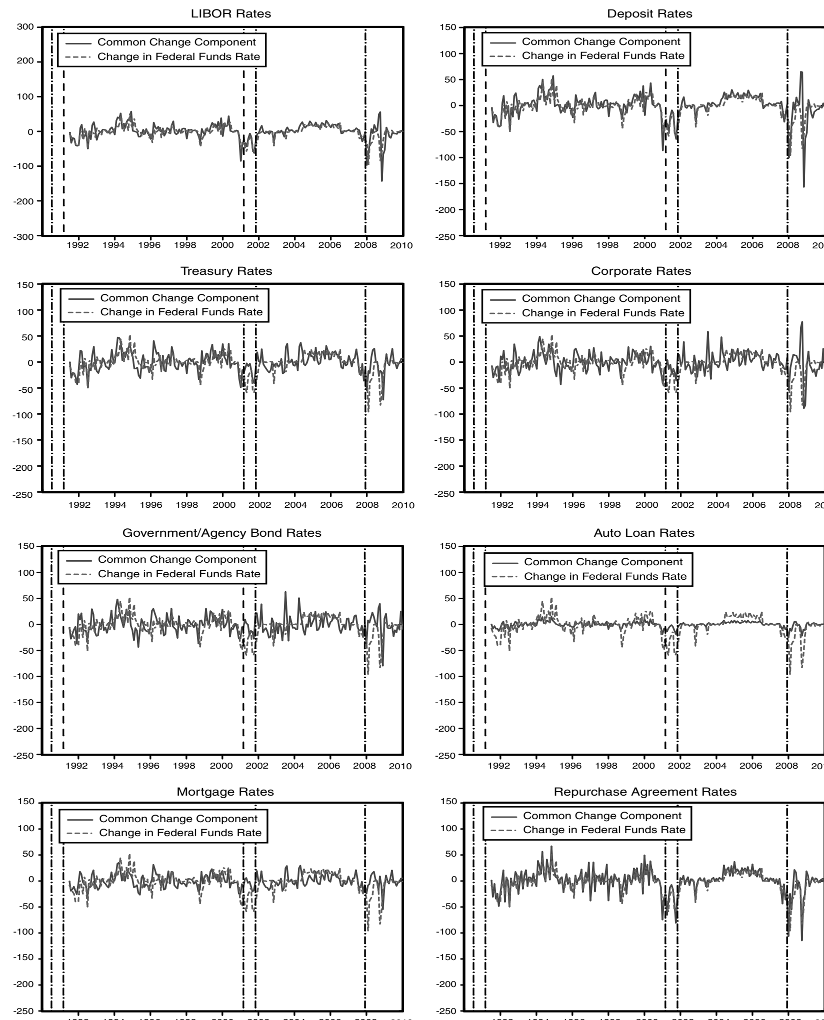
One should recognize that in each credit market category (defined by weights, w), changes in interest rates that stem from sources that are common across all credit markets, $w\Lambda_k F_{k,t}$, will not necessarily correspond to the behavior of average changes in these rates, wX_t. This is shown, for example, in Figure 3 where the difference between the common component of auto loan rate changes and average auto loan rate changes is evident. More important, having extracted the component of each rate change that is related to common sources, Figure 4 plots these common change components against changes in the effective federal funds rate for each of the eight broad credit markets defined above. It is apparent that the different components capturing the effects

Figure 3 Common and Average Rate Variations in Selected Credit Markets



of common forces look different across various credit market segments. However, the volatility of these change components tends to be similar to that of the effective federal funds rate. The question then is: What does the distribution of correlations between the different common change components in interest rates and changes in the effective federal funds rate look like? As mentioned earlier, changes in the effective federal funds rate may constitute only a rough summary of changes in monetary policy. A more general approach might be to examine a projection of common changes across individual interest rates on different aspects of changes in monetary policy, although ultimately not all relevant aspects of monetary policy are easily quantifiable or measured. For now, however, we focus on the effective federal funds rate.

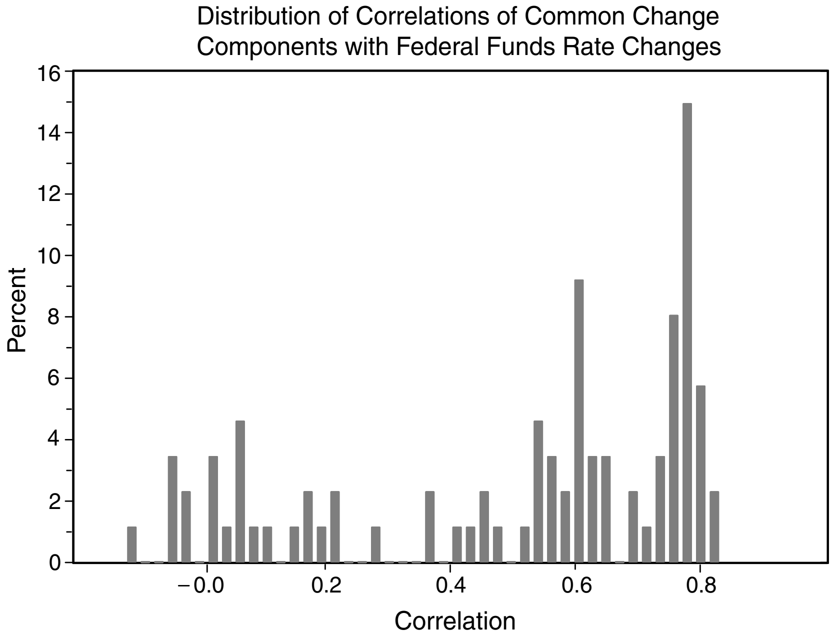
Figure 4 Common Rate Change Components and Federal Funds Rate Changes



Co-movement in Interest Rate Changes and the Federal Funds Rate

Figure 5 shows the histogram of the correlations between common change components in each interest rate, $\Lambda_k^i F_{k,t}$, and changes in the federal funds rate. While some of the common change components in interest rates seem highly correlated with changes in the federal funds rate, there are also many

Figure 5 Changes in the Federal Funds Rate and Common Changes in Interest Rates



other interest rates for which that is not the case. The median correlation in this case is 0.60 while the mean is 0.50. Table 3A provides a ranking of correlations across the eight credit market segments examined in this article.[3]

Interestingly, the common change components least correlated with changes in the federal funds rate are found in the government and agency bond and corporate credit markets. This finding may be interpreted in mainly two ways. First, although the common change components play an important role in driving corporate rates in Table 2A, these components likely reflect aggregate disturbances (or internal co-movement) that are somewhat unrelated to monetary policy. Second, to the extent that these rates include longer-term rates, they reflect monetary policy more indirectly through changes in expected future short rates. For example, changes in beliefs regarding future productivity will likely affect the perceived path of future federal funds rates. In contrast, we also see in Table 3A that the common change components in deposit and LIBOR rates are relatively highly correlated with changes in

[3] A listing of all correlations between $\Lambda^i_k F_{k,t}$ and Δr^{fed}_t can be found in the Appendix (Table 6).

Table 3 Correlation of Changes in Federal Funds Rate with Common Components by Interest Rate Category

Table 3A: Correlation of Common Components for Credit Markets with Changes in Federal Funds Rate

Series	**Correlation**
Government/Agency	0.131
Corporate	0.212
Mortgage	0.315
Treasury	0.501
Deposit	0.616
LIBOR	0.632
Repurchase Agreements	0.756
Auto	0.769

Table 3B: Correlation of Common Components for Treasury Securities with Changes in Federal Funds Rate

Series	**Correlation**
Three-Month Bill	0.776
Six-Month Bill	0.730
One-Year Bill	0.640
Two-Year Note	0.485
Three-Year Note	0.400
Five-Year Note	0.290
Seven-Year Note	0.220
10-Year Note	0.155
20-Year Bond	0.066
30-Year Bond	0.070

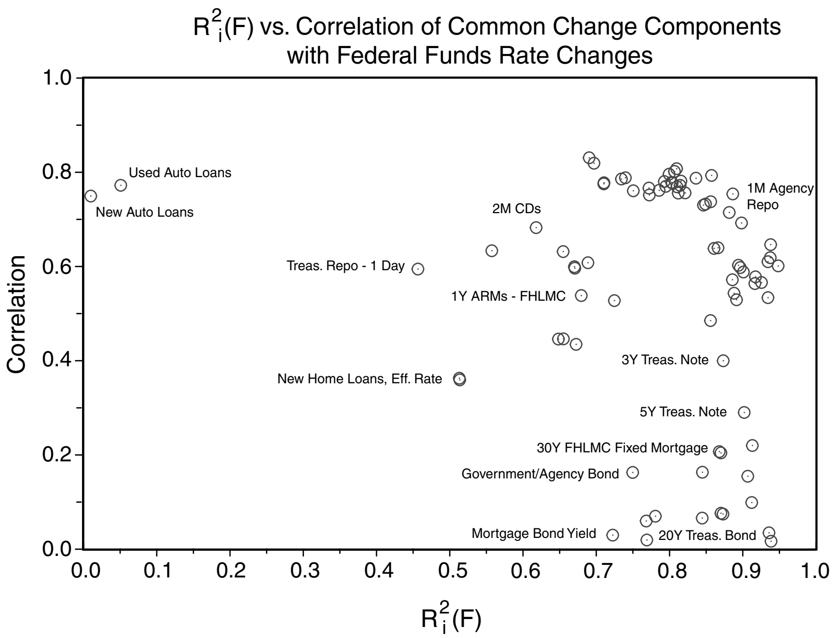
Notes: Monthly rates.

the federal funds rate. Moreover, Table 2A also suggests that the variations in these rates are, for the most part, accounted for by common sources of variations across interest rates. We conclude, therefore, that changes in monetary policy, as captured by changes in the federal funds rate, have played a fundamental role in driving deposit and LIBOR rates.

Table 3B provides the same statistics for Treasury rates of different maturities. As expected, the correlation between the common change component of Treasury bill rates and changes in the federal funds rate is decreasing in maturity, starting at 0.78 for the three-month rate and ending at 0.07 for the 30-year rate. Therefore, even if the common change component of 30-year rates plays a large role in explaining its variations (recall Table 2B), Table 3B is consistent with the conventional view that 30-year rates reflect other more fundamental aggregate changes in the economy rather than contemporaneous changes in policy.

Figure 6 summarizes the results thus far in the form of a scatter plot with $R_i^2(F)$ on the x-axis and $corr(\Lambda_k^i F_{k,t}, \Delta r_t^{\text{fed}})$ on the y-axis. A point near the

Figure 6 Effects of Monetary Policy Across Credit Markets



lower left-hand corner, where both statistics are near zero, would indicate that changes in interest rates are entirely disconnected from changes in the federal funds rate and, in essence, driven by more idiosyncratic considerations. The opposite is true near the top right-hand corner where both statistics are close to 1. Interestingly, the common components for auto loan rates have high correlations with changes in the federal funds rate, so that the common variation in these rates seems related to changes in monetary policy to a nontrivial extent, but also have extremely low $R_i^2(F)$. Put another way, although the common variation in auto loan rates is related to changes in the federal funds rate, their overall variation is ultimately driven by idiosyncratic considerations. There are also several rates in the lower right-hand corner of the plot. Variation in these rates is explained almost entirely by the common variation. However, the common components for these rates appear disconnected from monetary policy, as defined by the federal funds rate. Some of these rates include corporate bonds, fixed-rate mortgages, and long-term Treasury notes and bonds, and all of them have maturities of at least five years. Finally, Figure 6 also includes several rates near the top right-hand corner of the graph, namely several deposit, repurchase agreement, and Treasury bill rates, in which

Table 4 Correlation of Changes in Federal Funds Rate with Common Components by Interest Rate Category Over Different Sample Periods

Table 4A: Correlation of Common Components for Credit Markets with Changes in Federal Funds Rate

	Correlation	
Credit Market Segment	**1991:7–2001:2**	**2001:3–2009:12**
Government/Agency	0.26	0.03
Corporate	0.33	0.14
Mortgage	0.41	0.24
Treasury	0.55	0.47
LIBOR	0.67	0.60
Deposit	0.67	0.58
Repurchase Agreements	0.72	0.79
Auto	0.75	0.79

Table 4B: Correlation of Common Components for Treasury Securities with Changes in Federal Funds Rate

	Correlation	
Treasury Security Maturity	**1991:7–2001:2**	**2001:3–2009:12**
Three-Month Bill	0.76	0.80
Six-Month Bill	0.72	0.75
One-Year Bill	0.65	0.64
Two-Year Note	0.54	0.45
Three-Year Note	0.48	0.34
Five-Year Note	0.40	0.20
Seven-Year Note	0.35	0.12
10-Year Note	0.30	0.04
20-Year Bond	0.22	−0.06
30-Year Bond	0.23	−0.06

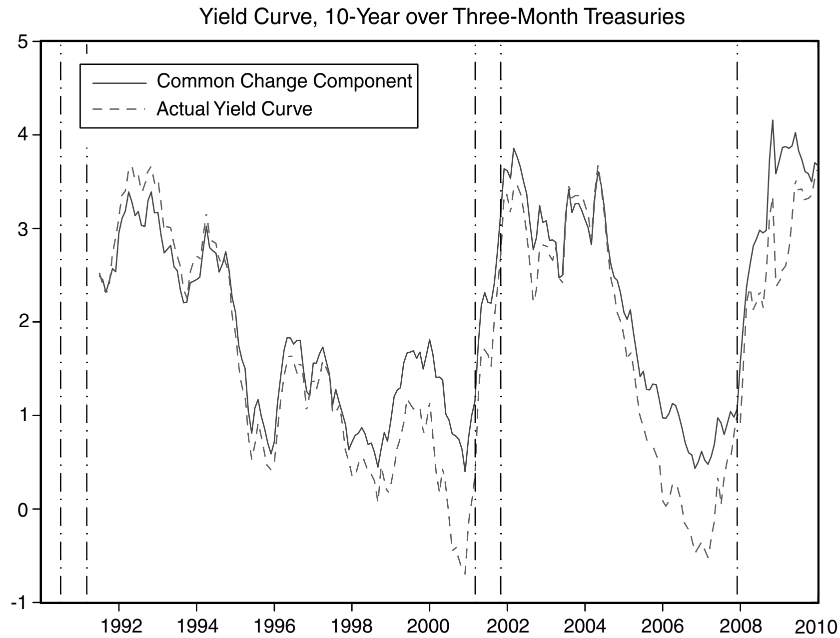
Notes: Monthly rates.

variations therefore appear closely related to changes in contemporaneous monetary policy.

Robustness Across Different Sample Periods

To analyze if this behavior has changed over time, we split the data into two subsamples: July 1991–February 2001 and March 2001–December 2009. We then calculate the correlations of the common components with changes in the effective federal funds rate over these two periods. We chose the breakpoint to be February 2001 to keep the subsamples roughly the same size, and because this is the month prior to the National Bureau of Economic Research peak of the 2001 recession. Table 4A shows the correlations for the eight broad groups

Figure 7 Common Changes in Interest Rates and the Yield Curve



described previously for each subsample. The ordering of the correlations for each credit market is essentially the same across the two periods. The most noticeable differences are seen in the mortgage, corporate, and government and agency bond markets. For these three groups, the correlations are moderately higher in the first subsample, indicating that disturbances less directly related to the contemporaneous federal funds rate have become more important in explaining common variation in these interest rate changes over time. This finding runs somewhat counter to the view in Taylor (2007) that an easy monetary policy kept long-term interest rates too low, thereby contributing to the housing boom. Rather, it is more consistent with the emphasis given by Bernanke (2010) to the role of other factors in keeping long-term interest low during the early 2000s.

The only two groups that saw an increase in correlations over the two periods are auto rates and repurchase agreement rates, though the increases are relatively small. Table 4B shows the analogous correlations for the common components of individual Treasury rates of different maturities. Interestingly, short-term Treasury bill rates have similar correlations across the two periods. However, at longer maturities, correlations between the common components and the federal funds rate have decreased in the later period, with the correlations for 20-year and 30-year Treasury bonds becoming slightly negative in the recent subsample.

As a final examination, we plot the common change component of the yield curve against the yield curve calculated from the raw data. These are shown in Figure 7. We define the yield curve as the 10-year Treasury note yield less the three-month Treasury bill yield. The main periods in which the two series deviate from each other are at their relative peaks and troughs, in particular in 1992, 2000, and 2006. However, overall the two series co-move strongly together, indicating that much of the spreads in rates of different maturities over time resides in how common shocks affect those rates rather than more idiosyncratic considerations.

5. CONCLUDING REMARKS

In this article, we use principal component methods to assess the importance of changes in the federal funds rate in driving interest rate changes across a variety of credit markets. Our findings suggest that most of the variability in interest rate changes across these markets can be explained by a small number of common components. In particular, four components explain approximately 80 percent of the total variation in interest rate changes. One notable exception is the auto loan market, in which interest rate variation is almost entirely idiosyncratic.

For most of our sample, the common variation in interest rate changes is relatively highly correlated with federal funds rate changes. This suggests that common movements in interest rates to a large extent reflect changes in monetary policy rather than other aggregate disturbances. That said, there nevertheless remains a moderate number of rates for which the common components, while explaining a significant portion of their variability, are not highly correlated with the federal funds rate. Therefore, these rates, which include mainly those with longer maturities such as mortgage rates, are driven to a greater extent by aggregate forces other than short-term changes in monetary policy. Finally, the analysis also suggests that movements in the auto loan market are almost entirely driven by idiosyncratic considerations rather than changes in the federal funds rate.

APPENDIX

This appendix shows that the solution to the principal component problem (3) also solves the least square problem described in (7). In particular, combining equations (7) and (8) gives

$$\min_{\{f_1,\ldots,f_k\}_{t=1}^{T},\Lambda_k} T^{-1}\sum_{t=1}^{T}(X_t - \Lambda_k F_{k,t})'(X_t - \Lambda_k F_{k,t}) \text{ subject to } \Lambda_k'\Lambda_k = I_k. \tag{16}$$

Suppose that Λ_k were known. Then the solution for $F_{k,t}$ would simply be given by the standard least square formula,

$$F_{k,t}(\Lambda_k) = (\Lambda_k'\Lambda_k)^{-1}\Lambda_k' X_t.$$

Substituting this solution into (16) yields

$$\min_{\Lambda_k} T^{-1}\sum_{t=1}^{T} X_t'[I_k - \Lambda_k(\Lambda_k'\Lambda_k)\Lambda_k']X_t,$$

or equivalently,

$$\max_{\Lambda_k} T^{-1}\sum_{t=1}^{T} X_t'\Lambda_k(\Lambda_k'\Lambda_k)\Lambda_k' X_t.$$

Now, note that this last expression is a scalar. Hence, we can re-write the problem as

$$\max_{\Lambda_k} tr\left(T^{-1}\sum_{t=1}^{T} X_t'\Lambda_k(\Lambda_k'\Lambda_k)\Lambda_k' X_t\right),$$

or

$$\max_{\Lambda_k} T^{-1}\sum_{t=1}^{T} tr\left(X_t'\Lambda_k(\Lambda_k'\Lambda_k)\Lambda_k' X_t\right).$$

Using the properties of the trace operator, this last expression can also be expressed as

$$\max_{\Lambda_k} T^{-1}\sum_{t=1}^{T} tr\left((\Lambda_k'\Lambda_k)^{-1/2}\Lambda_k'(X_t X_t')\Lambda_k(\Lambda_k'\Lambda_k)^{-1/2}\right),$$

or

$$\max_{\Lambda_k} tr\left((\Lambda_k'\Lambda_k)^{-1/2}\Lambda_k' T^{-1}\sum_{t=1}^{T}(X_t X_t')(\Lambda_k'\Lambda_k)^{-1/2}\Lambda_k(\Lambda_k'\Lambda_k)^{-1/2}\right).$$

Given the notation introduced in the text, one can observe that $T^{-1}\sum_{t=1}^{T}(X_t X_t')$ is simply $XX' = \Sigma_{XX}$. It follows that the least-square problem defined in (16) is equivalent to solving $\max_{\Lambda_k}\ \Lambda_k'\Sigma_{XX}\Lambda_k$ subject to $\Lambda_k'\Lambda_k = I_k$.

Table 5 Monthly Changes (in Basis Points)

Rate	Mean	Std. Dev.	Skewness	Kurtosis	Min	Max
Federal Funds [Effective] Rate	−2.6	19.9	−1.3	6.5	−96	53
One-Month London Interbank Offer Rate: Based on U.S.$	−2.6	28.0	−2.5	19.8	−219	88
Three-Month London Interbank Offer Rate: Based on U.S.$	−2.7	25.2	−1.9	15.2	−178	94
Six-Month London Interbank Offer Rate: Based on U.S.$	−2.7	23.8	−1.2	7.5	−122	65
One-Year London Interbank Offer Rate: Based on U.S.$	−2.7	25.8	−0.5	4.5	−98	74
One-Month Certificates of Deposit, Secondary Market	−2.6	29.2	−2.8	23.7	−241	98
Two-Month Certificate of Deposit	−2.6	29.1	−1.6	12.9	−172	119
Three-Month Certificates of Deposit, Secondary Market	−2.6	26.4	−2.2	17.4	−196	80
Six-Month Certificates of Deposit, Secondary Market	−2.7	25.7	−1.5	9.8	−154	71
Nine-Month Certificate of Deposit	−2.7	29.8	−1.4	11.0	−164	124
One-Month Eurodollar Deposits, London Bid	−2.5	32.3	−3.0	30.6	−285	140
Three-Month Eurodollar Deposits, London Bid	−2.5	28.8	−1.9	20.5	−220	136
Six-Month Eurodollar Deposits, London Bid	−2.6	26.5	−0.9	9.0	−143	109
U.S. Dollar: Eurocurrency Rate, Short-Term	−2.6	20.4	−1.6	8.0	−103	43
U.S. Dollar: Seven-Day Eurocurrency Rate	−2.5	24.9	−3.8	33.2	−229	55
U.S. Dollar: One-Month Eurocurrency Rate	−2.6	29.1	−2.2	19.9	−225	125
U.S. Dollar: Three-Month Eurocurrency Rate	−2.6	26.7	−1.1	11.6	−152	129
U.S. Dollar: Six-Month Eurocurrency Rate	−2.6	26.2	−0.9	8.2	−116	109
U.S. Dollar: One-Year Eurocurrency Rate	−2.7	27.3	−0.5	5.2	−106	98
Three-Month Treasury Bills, Secondary Market	−2.5	21.2	−1.1	5.0	−86	46
Six-Month Treasury Bills, Secondary Market	−2.5	21.0	−0.7	4.2	−73	51
Three-Month Treasury Bills	−2.5	20.8	−1.0	4.6	−83	45
Six-Month Treasury Bills	−2.5	20.8	−0.7	4.2	−75	52
Three-Month Treasury Bill Market Bid Yield at Constant Maturity	−2.6	22.0	−1.1	5.0	−89	49
Six-Month Treasury Bill Market Bid Yield at Constant Maturity	−2.6	22.2	−0.7	4.1	−77	54
One-Year Treasury Bill Yield at Constant Maturity	−2.7	23.4	−0.4	3.7	−79	60

Table 5 (Continued) Monthly Changes (in Basis Points)

Rate	Mean	Std. Dev.	Skewness	Kurtosis	Min	Max
Two-Year Treasury Note Yield at Constant Maturity	−2.7	26.1	0.0	2.9	−69	63
Three-Year Treasury Note Yield at Constant Maturity	−2.7	26.9	0.1	2.8	−69	65
Five-Year Treasury Note Yield at Constant Maturity	−2.5	26.1	0.1	2.9	−77	60
Seven-Year Treasury Note Yield at Constant Maturity	−2.2	24.6	0.1	3.3	−93	61
10-Year Treasury Note Yield at Constant Maturity	−2.0	23.4	−0.1	4.4	−111	65
20-Year Treasury Bond Yield at Constant Maturity	−1.7	20.8	−0.2	5.4	−109	58
30-Year Treasury Bond Yield at Constant Maturity	−1.7	19.9	−0.4	6.0	−110	51
Long-Term Treasury Composite, Over 10 Years	−1.9	20.8	−0.1	5.4	−109	59
One-Month Nonfinancial Commercial Paper	−2.7	22.2	−1.1	6.1	−94	68
Three-Month Nonfinancial Commercial Paper	−2.7	21.2	−1.0	6.1	−98	56
One-Month Financial Commercial Paper	−2.6	23.3	−1.7	10.8	−148	64
Three-Month Financial Commercial Paper	−2.6	23.1	−2.2	15.4	−165	61
Moody’s Seasoned Aaa Corporate Bond Yield	−1.7	18.4	−0.3	7.6	−107	63
Moody’s Seasoned Baa Corporate Bond Yield	−1.7	21.7	1.6	15.0	−76	157
Citigroup Global Markets: U.S. Broad Investment Grade Bond Yield	−2.4	28.6	0.3	4.6	−89	118
Citigroup Global Markets: Credit (Corporate) Bond Yield	−2.3	27.6	0.5	6.7	−116	124
Citigroup Global Markets: Credit (Corporate) Bond Yield: AAA/AA	−2.8	28.6	−0.7	9.7	−179	112
Citigroup Global Markets: Credit (Corporate) Bond Yield: AAA/AA 10+ Years	−1.6	23.8	−0.1	7.0	−115	90
Citigroup Global Markets: Credit (Corporate) Bond Yield: A	−2.3	28.1	0.4	6.8	−127	133
Citigroup Global Markets: Credit (Corporate) Bond Yield: BBB	−2.1	30.1	1.8	16.0	−80	222
Citigroup Global Markets: Credit (Corporate) Bond Yield: Finance	−1.9	33.6	1.1	12.9	−130	227
Citigroup Global Markets: Credit (Corporate) Bond Yield: Utility	−2.2	30.1	0.7	12.0	−147	188
Citigroup Global Markets: Credit (Corporate) Bond Yield: Industrial	−2.3	27.3	1.3	11.3	−80	181
Citigroup Global Markets: Gov’t Sponsored Bond Yield, U.S. Agency/Supranational	−2.7	24.7	−0.1	3.8	−86	77
Citigroup Global Markets: Gov’t Agency Bond Yield	−2.5	26.2	0.0	3.2	−85	69
Bond Buyer Index: State/Local Bonds, 20-Year, Genl Obligation	−1.3	16.0	0.7	4.6	−49	64
Auto Finance Company Interest Rates: New Car Loans	−4.2	65.2	−1.1	9.4	−392	172
Auto Finance Company Interest Rates: Used Car Loans	−3.3	21.2	0.3	3.8	−59	74
Citigroup Global Markets: Mortgage Bond Yield	−2.5	37.9	0.9	8.3	−110	200
Home Mortgage Loans: Effective Rate, All Loans Closed	−1.9	13.2	−0.3	4.7	−63	34

Table 5 (Continued) Monthly Changes (in Basis Points)

Rate	Mean	Std. Dev.	Skewness	Kurtosis	Min	Max
Purchase of Newly Built Homes: Effective Rate, All Loans	−2.0	14.2	−0.3	4.2	−56	35
Purchase of Previously Occupied Homes: Effective Rate, All Loans	−1.9	13.7	−0.3	4.8	−67	35
Contract Rates on Commitments:						
Conventional 30-Year Mortgages, FHLMC	−2.1	20.6	0.5	4.0	−76	64
Purchase of New Single-Family Home: Contract Interest Rate	−1.9	13.7	−0.3	4.3	−55	35
Purchase of Existing Single-Family Home: Contract Interest Rate	−1.8	13.3	−0.4	5.2	−67	36
FHLMC: 30-Year Fixed-Rate Mortgages: U.S.	−2.1	20.6	0.6	4.0	−76	64
FHLMC: 1-Year Adjustable Rate Mortgages: U.S.	−1.3	14.5	0.6	4.3	−39	56
Treasury Repo - One Day	−2.5	43.1	−0.5	5.6	−165	168
Treasury Repo - One Week	−2.6	25.2	−1.5	9.7	−140	69
Treasury Repo - One Month	−2.6	23.3	−2.4	13.2	−140	59
Treasury Repo - Three Months	−2.6	23.6	−2.2	12.0	−135	48
Treasury Reverse Repo - One Day	−2.5	43.3	−0.4	5.6	−180	168
Treasury Reverse Repo - One Week	−2.6	27.1	−1.1	7.7	−130	82
Treasury Reverse Repo - One Month	−2.6	23.4	−2.2	11.8	−135	58
Treasury Reverse Repo - Three Months	−2.6	22.8	−1.9	10.3	−125	48
MBS Repo - One Day	−2.6	43.0	−0.3	5.1	−140	168
MBS Repo - One Week	−2.6	25.5	−0.8	6.5	−110	85
MBS Repo - One Month	−2.6	22.9	−1.9	9.7	−117	62
MBS Repo - Three Months	−2.6	22.8	−1.9	10.3	−116	62
MBS Reverse Repo - One Day	−2.5	44.2	−0.4	5.1	−153	168
MBS Reverse Repo - One Week	−2.7	29.1	−1.0	6.9	−120	90
MBS Reverse Repo - One Month	−2.6	23.5	−1.8	9.4	−120	60
MBS Reverse Repo - Three Months	−2.6	23.0	−1.9	10.3	−120	56
Agency Repo - One Day	−2.5	39.8	−0.3	6.0	−145	168
Agency Repo - One Week	−2.6	24.6	−1.1	6.9	−115	72
Agency Repo - One Month	−2.6	22.9	−2.2	11.6	−130	59
Agency Repo - Three Months	−2.6	23.4	−2.2	11.3	−125	43
Agency Reverse Repo - One Day	−2.5	43.5	−0.7	7.3	−225	168
Agency Reverse Repo - One Week	−2.6	27.0	−1.0	6.8	−115	82
Agency Reverse Repo - One Month	−2.6	24.0	−2.2	11.8	−130	58
Agency Reverse Repo - Three Months	−2.6	23.9	−2.2	10.8	−125	41

Table 6 R-Squared and Correlation of Factor Components with Federal Funds Rate (Monthly Data)

Rate	R^2	Correlation
Auto Finance Company Interest Rates: New Car Loans	0.010	0.750
Auto Finance Company Interest Rates: Used Car Loans	0.051	0.772
Treasury Repo - One Day	0.457	0.595
Bond Buyer Index: State/Local Bonds, 20-Year, Genl Obligation	0.459	−0.027
Purchase of New Single-Family Home: Contract Interest Rate	0.513	0.363
Purchase of Newly Built Homes: Effective Rate, All Loans	0.514	0.359
Treasury Reverse Repo - One Day	0.558	0.634
Two-Month Certificate of Deposit	0.618	0.683
Moody's Seasoned Baa Corporate Bond Yield	0.643	−0.118
Purchase of Existing Single-Family Home: Contract Interest Rate	0.648	0.446
Agency Repo - One Day	0.655	0.632
Purchase of Previously Occupied Homes: Effective Rate, All Loans	0.655	0.447
MBS Repo - One Day	0.670	0.600
MBS Reverse Repo - One Day	0.671	0.597
Home Mortgage Loans: Effective Rate, All Loans Closed	0.672	0.435
FHLMC: One-Year Adjustable Rate Mortgages: U.S.	0.680	0.538
Agency Reverse Repo - One Day	0.689	0.608
U.S. Dollar: Eurocurrency Rate, Short-Term	0.697	0.819
Three-Month Treasury Bill Market Bid Yield at Constant Maturity	0.710	0.776
Three-Month Treasury Bills, Secondary Market	0.711	0.778
Citigroup Global Markets: Mortgage Bond Yield	0.722	0.030
Nine-Month Certificate of Deposit	0.725	0.528
Three-Month Treasury Bills	0.734	0.786
Agency Repo - One Week	0.740	0.789
Citigroup Global Markets: Gov't Agency Bond Yield	0.750	0.163

Table 6 (Continued) R-Squared and Correlation of Factor Components with Federal Funds Rate (Monthly Data)

Rate	R^2	Correlation
MBS Reverse Repo - One Week	0.751	0.761
Citigroup Global Markets: Credit (Corporate) Bond Yield: Utility	0.764	−0.037
Citigroup Global Markets: Credit (Corporate) Bond Yield: Finance	0.768	0.060
Moody’s Seasoned Aaa Corporate Bond Yield	0.769	0.020
Treasury Repo - Three Months	0.772	0.766
Treasury Reverse Repo - One Week	0.772	0.752
30-Year Treasury Bond Yield at Constant Maturity	0.781	0.070
Agency Reverse Repo - One Week	0.786	0.762
MBS Repo - One Week	0.793	0.781
Citigroup Global Markets: Credit (Corporate) Bond Yield: BBB	0.793	−0.053
Treasury Reverse Repo - Three Months	0.795	0.770
One-Month Nonfinancial Commercial Paper	0.799	0.796
Agency Repo - Three Months	0.803	0.779
Treasury Repo - One Month	0.804	0.777
MBS Reverse Repo - One Month	0.807	0.802
MBS Repo - One Month	0.810	0.808
Agency Reverse Repo - Three Months	0.810	0.769
MBS Reverse Repo - Three Months	0.812	0.756
MBS Repo - Three Months	0.815	0.772
Treasury Reverse Repo - One Month	0.815	0.781
Treasury Repo - One Week	0.821	0.756
Citigroup Global Markets: Credit (Corporate) Bond Yield: AAA/AA 10+ Years	0.833	−0.042
Agency Reverse Repo - One Month	0.836	0.788
20-Year Treasury Bond Yield at Constant Maturity	0.844	0.066
Citigroup Global Markets: Gov’t Sponsored Bond Yield: U.S. Agency/Supranational	0.845	0.163
Six-Month Treasury Bill Market Bid Yield at Constant Maturity	0.846	0.730
Six-Month Treasury Bills, Secondary Market	0.849	0.733
Two-Year Treasury Note Yield at Constant Maturity	0.856	0.485
Six-Month Treasury Bills	0.856	0.738
Agency Repo - One Month	0.857	0.793
U.S. Dollar: Seven-Day Eurocurrency Rate	0.861	0.638

Table 6 (Continued) R-Squared and Correlation of Factor Components with Federal Funds Rate (Monthly Data)

Rate	R^2	Correlation
One-Year Treasury Bill Yield at Constant Maturity	0.866	0.640
FHLMC: 30-Year Fixed-Rate Mortgages: U.S.	0.868	0.207
Long-Term Treasury Composite, Over 10 Years	0.870	0.076
Contract Rates on Commitments: Conventional 30-Yr Mortgages, FHLMC	0.870	0.204
Citigroup Global Markets: U.S. Broad Investment Grade Bond Yield	0.873	0.074
Three-Year Treasury Note Yield at Constant Maturity	0.873	0.400
One-Month Financial Commercial Paper	0.882	0.715
U.S. Dollar: One-Month Eurocurrency Rate	0.886	0.572
Three-Month Nonfinancial Paper	0.886	0.754
One-Month Eurodollar Deposits, London Bid	0.888	0.543
Citigroup Global Markets: Credit (Corporate) Bond Yield: Industrial	0.889	−0.016
U.S. Dollar: One-Year Eurocurrency Rate	0.892	0.529
One-Month London Interbank Offer Rate: Based on U.S.$	0.894	0.603
One-Month Certificates of Deposit, Secondary Market	0.897	0.598
Three-Month Financial Commercial Paper	0.898	0.693
One-Year London Interbank Offer Rate: Based on U.S.$	0.901	0.588
Five-Year Treasury Note Yield at Constant Maturity	0.902	0.290
10-Year Treasury Note Yield at Constant Maturity	0.907	0.155
Citigroup Global Markets: Credit (Corporate) Bond Yield: AAA/AA	0.912	0.099
Seven-Year Treasury Note Yield at Constant Maturity	0.913	0.220
U.S. Dollar: Three-Month Eurocurrency Rate	0.916	0.564
U.S. Dollar: Six-Month Eurocurrency Rate	0.917	0.578
Six-Month Eurodollar Deposits, London Bid	0.926	0.566
Three-Month London Interbank Offer Rate: Based on U.S.$	0.934	0.610
Three-Month Eurodollar Deposits, London Bid	0.934	0.534
Citigroup Global Markets: Credit (Corporate) Bond Yield: A	0.936	0.035
Six-Month Certificates of Deposit, Secondary Market	0.937	0.619
Six-Month London Interbank Offer Rate: Based on U.S.$	0.938	0.646
Citigroup Global Markets: Credit (Corporate) Bond Yield	0.939	0.017
Three-Month Certificates of Deposit, Secondary Market	0.948	0.601

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