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Did Prices Really Soar after the Euro Cash Changeover? Evidence from ATM Withdrawals*

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The introduction of the euro notes and coins in the first two months of 2002 was followed by a lively debate on the alleged inflationary effects of the new currency. In Italy, as in the rest of the euro area, survey-based measures signaled a much sharper rise in inflation than that measured by the official price indices, the quality of which was called into question. In this paper we gather indirect evidence on the behavior of prices from the analysis of cash withdrawals from automated teller machine (ATM) terminals. Since these data do not rely on official inflation statistics, they provide an independent check for the latter. We present a simple set of assumptions to test the hypothesis that, after the introduction of the euro notes and coins, consumer prices increased more than was recorded by the official statistics. We do not find evidence in support of this hypothesis.

JEL Codes: E31, E41, E50.

Whatever the experts say, many European consumers still feel retailers are masking price increases with the changeover to the euro.
(*Wall Street Journal Europe*, January 28, 2002)

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Two out of three eurozone consumers felt they were ripped off by retailers during the changeover to pricing in euros, according to the European Commission.... Germany, France and Netherlands were the countries with the highest percentage of people feeling cheated.

(*Financial Times*, March 1, 2002)

German consumers dubbed the currency the Teuro (teuer is German for expensive).... Some consumers believe higher prices were the result of retailers rounding up prices as they switched out of their old national currencies into the euro. However, EU statisticians insisted prices had not been affected.

(*Financial Times*, December 12, 2002)

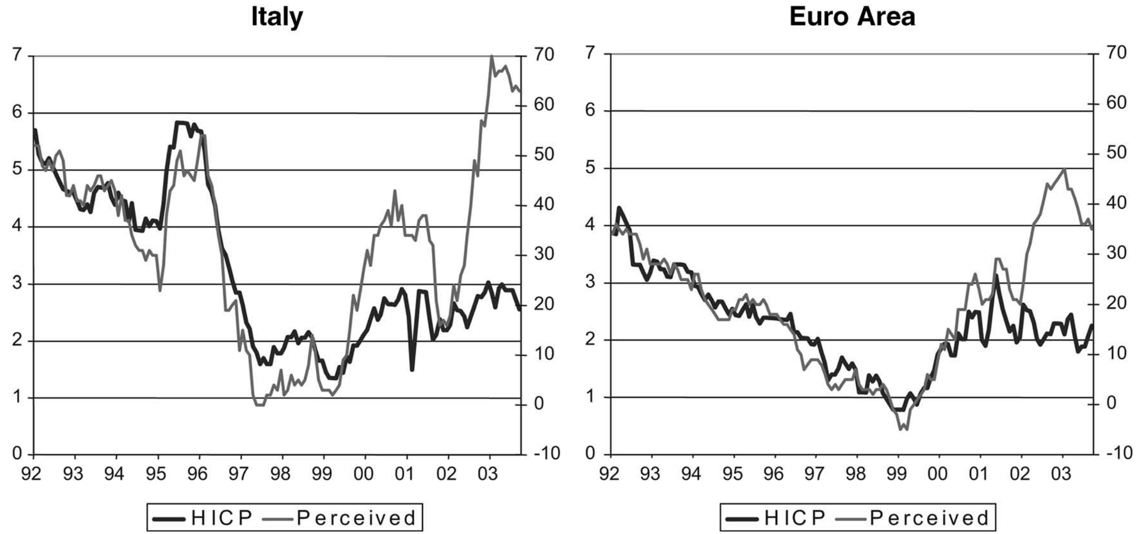
1. Introduction

There is a widespread perception among the citizens of the euro area that the introduction of the euro notes and coins in the first months of 2002 spurred a rise in prices that was much sharper than that measured by the national statistical offices (see European Central Bank 2002b, 2003a, 2003b). This phenomenon, illustrated in figure 1, has been the subject of countless newspaper articles and several official speeches by policymakers and politicians. While the quotations reported above refer to 2002, the perception that the euro brought about price increases is still vivid.[1]

It is somewhat puzzling that a change in the unit of account might have an impact on inflation. Indeed, a number of conjectures have been formulated to explain the discrepancy between inflation perceptions and the official statistics, emphasizing the role of psychological factors (see, e.g., Traut-Mattausch et al. 2004) and/or

[1]In May 2002 Professor Otmar Issing gave a speech in Mainz titled "The Euro—A Stable Currency for Europe." After the speech, the first question from the audience was about the "teuro" phenomenon. Seeing the look of disbelief with which his explanation was met, Issing replied, "You seem not to believe me. And even my wife doesn't believe me." This statement found wide coverage in the German press. Two years later, President Trichet still deemed it necessary to reassure European customers on this issue: "European citizens who still perceive that inflation is higher than measured by official indices should be assured that the official measures are accurate and that we will continue to maintain price stability in the future" (introductory statement after the European Central Bank's Governing Council meeting of April 2004).

Figure 1. Perceived and Official Inflation in Selected Euro-Area Countries (Monthly Data)



Source: Eurostat, European Commission.

Note: "HICP" (left axis) stands for inflation measured by the twelve-month growth rate of the Harmonized Index of Consumer Prices. "Perceived" (right axis) stands for perceived inflation based on surveys by the Institute for Studies and Economic Analyses (ISAE) and by the European Commission, as reported in the monthly press release "Business and Consumer Survey Results." It is computed as the difference between the share of respondents reporting that prices "strongly increased" (weight 1) or "moderately increased" (weight 0.5) and the share of respondents reporting "stable" (weight 0.5) or "decreased" (weight 1) prices.

the disproportionate influence of a few industry prices on individual perceptions. Gaiotti and Lippi (2004) and Hobijn, Ravenna, and Tambalotti (2006) analyze the dynamics of restaurant prices and find evidence consistent with a price hike (mainly driven by a lumping of price revisions in an industry where price revisions are normally infrequent). Deutsche Bundesbank (2004) provides comparable evidence for some German services (restaurants, cinemas, dry cleaners, and hairdressers). Other papers argue that inflation perceptions are mainly affected by the prices of goods that are cheaper and more frequently purchased (Del Giovane and Sabbatini 2005; Ehrmann 2005). Dziuda and Mastrobuoni (2005) and Mastrobuoni (2004) present a model that rationalizes why such goods are the ones that actually record greater price increases. While useful, these studies do not provide a direct answer to whether the *general* price level was measured with error during the changeover. Rather, they maintain the assumption that official statistics are correct. The main obstacle faced by researchers interested in verifying that assumption is the absence of reliable alternative inflation measures. The thesis that price increases were much larger than the increase measured by the national statistical offices, suggested by the indicators of perceived inflation, remains mostly based on anecdotal evidence.

This paper investigates the dynamics of the general price level in Italy after the introduction of the euro notes and coins (the so-called cash changeover), at the beginning of January 2002, by using data on currency withdrawn from the automated teller machine (ATM) network. We believe that this inference is useful because it relies on data collected and assembled by central banks, with methodologies that are completely independent of those used by the national statistical offices to measure prices. The basic steps of our investigation can be summarized as follows. We set up a simple model of ATM withdrawals and estimate it *prior* to the changeover, when official statistics were arguably correct. We then present a set of assumptions under which the estimated demand equation for ATM withdrawals can be used to back out price-level dynamics from the observed nominal time series for ATM withdrawals and consumer expenditures. Specifically, we show that if a bias materialized in the official data on prices after the changeover, but not in those on cash withdrawals and consumption expenditures, then extending the estimation sample to the changeover period (2002–03) should cause a specific form of instability in the estimated coefficients, which can be captured econometrically. Formally, we test the null hypothesis that the increase in consumer prices is correctly measured by official statistics after the changeover. Both a price-level-bias and an inflation-bias hypothesis are formulated and tested. The analysis fails to find evidence consistent with the occurrence of a price hike after the changeover. This result cannot be attributed to lack of power of the test: a counterfactual exercise suggests that our methodology is sufficiently powerful to identify an inflation bias greater than 0.5 percentage point.

Several reasons motivate our focus on Italy. First, the country is broadly representative of the euro area in terms of the discrepancy between official and perceived inflation (figure 1). Also, quarterly data on cash withdrawals are available, whereas comparable data are available only at an annual frequency for other euro-area countries (to our knowledge). It is important to explain why we focus on the flow of currency withdrawn from the ATM circuit rather than on more traditional monetary aggregates. The stock of currency experienced a strong decline from the beginning of 2001 (apparently reflecting weak demand of bank notes as a store of wealth due to the approaching currency changeover) and an equally strong rebound thereafter.[2] Among the traditional monetary aggregates, M1 is strongly affected by the erratic behavior of currency. M2 and M3 are comparatively less affected, but they are typically less related to transactions; in addition, over the recent past, the dynamics of these aggregates has been influenced by portfolio reasons, as repeatedly stressed by the European Central Bank. By contrast, there is no obvious reason why ATM withdrawals—mainly driven by transactions demand—should have been affected by these same factors. The data in figure 2 (shown in section 3) broadly confirm this view: neither the average number nor the unit value of ATM withdrawals made in each quarter by a typical cardholder shows the discontinuity that clearly

[2]See, e.g., European Central Bank (2002a). As we argue in what follows, this decline is likely due to the attempt by currency holders to run down their cash inventories. These are mainly held in large-denomination notes and are to a significant (although not easily quantifiable) extent held outside the national borders and/or for gray/black economy purposes.

stands out for the traditional monetary aggregates around the changeover.

The paper is organized as follows. The next section presents a simple model of the demand for ATM withdrawals, which is used as a guideline in the empirical analysis of section 3. Section 4 concludes.

2. A Simple Model of Aggregate ATM Withdrawals

This section presents a model aimed at interpreting the evolution of aggregate ATM withdrawals. To match the aggregate data, we first focus on the choice concerning the withdrawals by a representative ATM cardholder and, next, present an aggregation to account for the growing number of cardholders over our estimation period.[3]

Let i index an agent who possesses an ATM card and E^i denote that agent's nominal consumption expenditure over a given time span. To pay for E^i, the agent can use cash C^i withdrawn from an ATM, bearing a cost r^c, or some other means of payment Q (e.g., point of sale [POS], credit card, or checks), the cost of which is denoted by r^q. The demand schedule for the agent's ATM withdrawals is of the following form: $C^i = \Phi(r^q/r^c)E^i$, where the function $\Phi(\cdot)$ is increasing and concave.[4] This demand function stipulates that the proportion of cash withdrawals over total nominal expenditure is decreasing in the relative cost between ATM cash and that of alternative means of payments.

To give empirical content to these costs, in the analysis that follows we assume that the cost of ATM cash, r^c, is increasing in the nominal interest rate R (the value of which determines the amount of forgone interest on deposits) and decreasing in the size of the ATM network, d^{ATM}. Moreover, we proxy the cost of alternative payment means, r^q, using a measure of development of the POS network,

[3]Based on the Survey of Households' Income and Wealth conducted by the Bank of Italy, between 1989 and 2004 the proportion of households owning an ATM card rose from 15 percent to 66 percent.

[4]This cash/credit choice can be thought of along the lines of Lucas and Stockey (1987).

denoted by d^Q.[5] The individual proportion of cash expenditures is thus hypothesized to depend on the diffusion of the ATM and POS network and the nominal interest rate:

$$C^i = \Phi(d^Q/r^c(d^{ATM}, R))E^i. \tag{1}$$

Let us now consider the aggregation problem. In order to bring equation (1) to the data, we need to relate the nationwide demand for ATM cash to aggregate consumption, as data on total expenditure of ATM cardholders are not available. Let n and $\bar{E}/P$ denote the number of ATM cardholders and their aggregate expenditure (deflated by the price level), respectively. Analogously, let N and E/P denote the population size and real aggregate expenditures. We postulate the following:

$$\bar{E}/P = (n/N)(E/P)^{\delta}. \tag{2}$$

This equation states that the growth in the aggregate expenditure of cardholders is proportional to the growth of the population fraction of cardholders and to the growth rate of aggregate expenditures, with a constant of proportionality that may differ from 1.

Aggregating equation (1) over all cardholders and replacing $\bar{E}$ using (2) yields an expression relating the real aggregate flow of ATM withdrawals to real aggregate expenditure:

$$C/P = (n/N)\Phi(d^Q/r^c(d^{ATM}, R))(E/P)^{\delta}. \tag{3}$$

Equation (3) summarizes the determinants of *aggregate* ATM cash withdrawals discussed thus far. C is increasing in ATM diffusion (d^{ATM}) and decreasing in the nominal interest rate (R) and in the ease of resorting to noncash payments (d^Q). The elasticity of aggregate cash withdrawals with respect to aggregate expenditure is given by δ. It is immediate that if δ is equal to 1, then the price level drops from equation (3) and no information about it can be retrieved from that equation. For all other values, instead, one can invert equation (3) and use the information on C and E to back out

[5] In Italy, our data source for the tests presented in the next section, it is not possible to get cash back at POS. This option, available to customers in several industrialized countries, would have made this proxy questionable.

the price level. In section 3.2 we discuss the functional form adopted in equation (2) and review the available empirical evidence on δ, which is key for our identification of price-level dynamics.

3. Empirical Evidence

This section begins by presenting some descriptive evidence on ATM withdrawals during the period 1993–2003. Using equation (3) as a guide for the empirical specification, we estimate a currency-expenditure equation and formally test for the presence of measurement error in the aggregate price level. We conclude the section by exploring the power of the statistical test and analyzing the robustness of the estimates (considering, e.g., parameter instability and potential endogeneity of the regressors).

3.1 A Preliminary Look at the Data

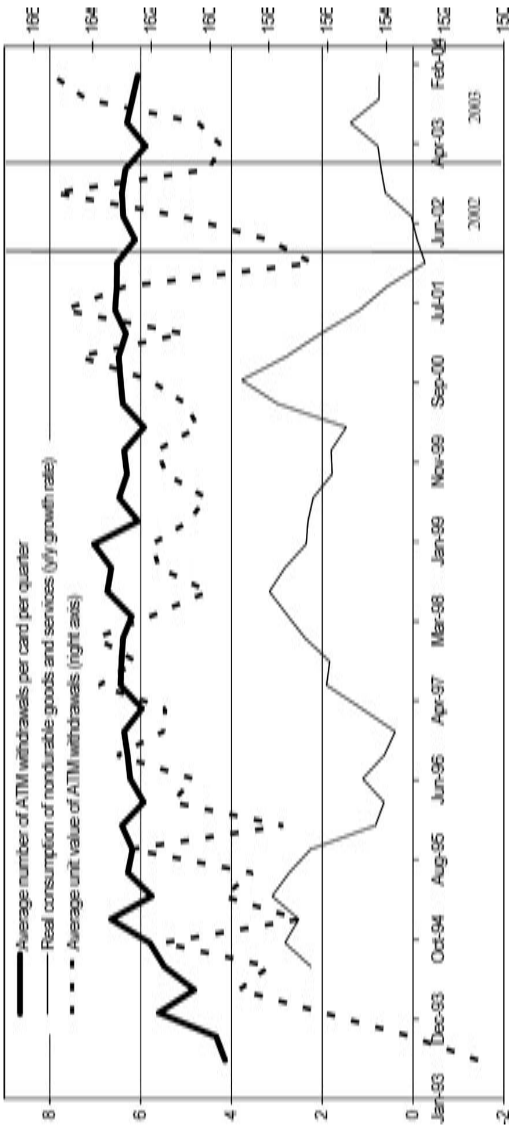
Our data set comprises quarterly time series over the 1993:Q2–2003:Q4 period.[6] Figure 2 shows that during the first three quarters of 2002, the average unit value of ATM withdrawal per card (dashed line) records a sharp increase (from €157 to €165), but only after an equally sharp fall in 2001.[7] Overall, the withdrawal per card after the changeover remains close to the values recorded in the previous five years.[8] The same conclusion holds for the frequency of

[6]We refer to two sources. The flow of cash withdrawn from ATMs in Italy, the number of ATM cards, the number of POS and ATM terminals, and the interest rate on checking accounts are provided to the Bank of Italy by the banking system for supervisory reasons. Data on consumption of services and nondurable goods, and the related deflators, are released by the Italian national statistical office (ISTAT). All the series used in the paper refer to Italy.

[7]This swing is likely due to the need by currency holders to run down their cash inventories. As mentioned in the introduction, a massive drop in currency outstanding was recorded in 2001. As the cash changeover was largely anticipated by the public, it seems plausible that, in the final months of the year, currency holders avoided withdrawing cash and used their inventory for day-to-day purchases. After the changeover, the average use of ATM cash returned to normal.

[8]Such stationarity, in a period of moderate but positive inflation, might reflect the increasing use of cash substitutes; the regression analysis below supports this hypothesis.

Figure 2. ATM Usage and Consumption Growth (Quarterly Data)



Source: Bank of Italy; ISTAT.

ATM usage (thick solid line), roughly constant at 6.3 withdrawals per card per quarter since 1995. Figure 2 also displays the behavior of households' real consumption of nondurable goods and services in the period (thin solid line). Growth begins to slow down in the second half of 2000, bottoming out in the fourth quarter of 2001.

At least prima facie, none of these time series display a behavior that might signal a sharp increase in the price level after the introduction of the euro bank notes and coins. However, this descriptive evidence can be potentially misleading, for at least two reasons. The first relates to our measure of consumption. Assume that inflation perceptions are correct and that the official nondurable goods deflator underestimates inflation in 2002–03; then, real-consumption growth in figure 2 would be correspondingly overestimated. Second, the descriptive evidence does not take into account several structural changes that have occurred over the last fifteen years—most notably, the diffusion of ATM and POS terminals, and the associated increase in aggregate cash withdrawals from ATMs and in the use of noncash payments via debit cards. Such developments typically follow low-frequency trends, so in our view it is unlikely that they may have obscured the effect of the hypothesized price jump on the demand for cash and alternative payment instruments in 2002–03. However, the model of section 2 suggests that they affect the demand for ATM withdrawals and should thus be taken into account. This is done in the next subsection.

3.2 Inference from an Estimated Currency-Expenditure Equation

A log-linear version of equation (3) yields

$$c_t - p_t = \alpha + \beta d_t^{ATM} + \gamma d_t^{Q} + \delta(e_t - p_t) + \varepsilon r_t + \lambda n_t + \eta_t, \qquad (4)$$

where η_t is an error term with variance σ_η^2. We measure c_t by the (log) nominal value of nationwide quarterly withdrawals from ATM terminals; we measure n_t by the (log) number of outstanding ATM cards (at the end of the quarter). r_t is the (log) interest rate on checking accounts. The diffusion of the ATM network, d^{ATM}, is proxied by the ratio between the number of ATM terminals nationwide and the number of ATM cards. Similarly, the diffusion of alternative payment instruments, d^Q, is measured by the ratio between the number of POS terminals and n_t. Finally, e_t and p_t are proxied by the (log) aggregate nominal consumption of services and nondurable goods, and by its deflator, respectively.[9]

Can equation (4) help us shed light on the issue at the core of this paper? For the answer to be positive, we must assume that all of the time series appearing in (4) are measured with no error until the fourth quarter of 2001.[10] This assumption seems reasonable, since until the end of 2001 there was no argument about data quality. It implies that the coefficients of (4) will not be affected by measurement problems if estimated over the pre-changeover period. Concerning the post-changeover period, we assume that all variables appearing in (4) are measured correctly, with the possible exception of the price level. Specifically, we allow for the possibility that p_t, the true (log) price level, may suffer from measurement error and deviate from its observed counterpart, p_t^o.

As anticipated in section 3, a condition on the elasticity of cash withdrawals with respect to real consumption (δ) must also be satisfied. It is easy to see that if $\delta = 1$, the price level cancels from both sides of (4), and the equation becomes uninformative on the price-level dynamics. Thus, an estimated $\delta \neq 1$ is necessary (and sufficient) to back out the price level after 2001 from (4).

We begin by estimating (4) using data until 2001:Q4. The results are reported in table 1, column 1.[11] The coefficients are in line with the model suggestions: the diffusion of noncash forms of payment reduces cash withdrawals, while the diffusion of ATM

[9]Based on equation (3), the population size, N, should also appear among the regressors. We omit it because it was roughly constant over the estimation period.

[10]More precisely, we are assuming that the measurement errors present in official statistics (e.g., the well-known price-index bias) are not affected by the changeover. Under our alternative hypothesis, we allow the euro changeover to bring about a source of measurement error in price measures *additional* to those that are ordinarily present.

[11]Three quarterly dummies (not shown) are included among the explanatory variables to account for seasonal effects.

Table 1. Estimates of Equations (4) and (6)

		Equation (6) Estimated over 1993:Q2–2003:Q4			
		Test of Hypothesis (5a)		Test of Hypothesis (5b)	
	Equation (4) Estimated over 1993:Q2–2001:Q4 1	**Estimating Only Dummy Coefficient 2**	**Estimating All Coefficients 3**	**Estimating Only Dummy Coefficient 4**	**Estimating All Coefficients 5**
ATM Terminals Diffusion (d_t^{ATM})	2.38**	2.38	2.34**	2.38	2.43**
	5.90	Constrained	7.00	Constrained	6.80
POS Terminals Diffusion (d_t^Q)	−17.05**	−17.05	−17.00**	−17.05	−15.75*
	−5.00	Constrained	−6.90	Constrained	−5.70
Real Consumption (log; $e_t - p_t$)	2.82**	2.82	2.75**	2.82	2.57**
	3.60	Constrained	4.00	Constrained	3.80
Number of ATM Cards (log; n_t)	0.89**	0.89	0.91**	0.89	0.89**
	5.60	Constrained	6.50	Constrained	6.20
Interest Rate on Checking Accounts (log; r_t)	−7.1e–4	−7.1e–4	3.4e–3	−7.1e–4	6.9e–3
	0.00	Constrained	0.20	Constrained	0.30
Constant	−8.82	−8.82	−8.57	−8.82	−5.42
	−0.80	Constrained	−1.10	Constrained	−0.60
Dummy 2002:Q1–2003:Q4	—	−6.2e–3	−6.4e–3	—	—
	—	−0.60	−0.40	—	—
Dummy 2002:Q1–2003:Q4*Linear Trend	—	—	—	−2.0e–3	−3.6e–3
	—	—	—	−1.00	−1.10
F-test for $\sigma_\nu^2 = \sigma_\eta^2$ vs. $\sigma_\nu^2 > \sigma_\eta^2$ (p-value)	—	0.90	0.96	0.95	0.99
Number of Observations	35	43	43	43	43
R^2	0.98	—	0.99	—	0.99
DW	1.47	1.59	1.58	1.60	1.60

Note: The dependent variable is the (log of) aggregate cash withdrawals from ATM in real terms. σ_η^2 and σ_ν^2 denote, respectively, the variance of the error term in equation (6) before and after 2001:Q4. OLS estimates. Heteroskedasticity-robust t-statistics are reported below each coefficient. * and ** denote, respectively, 5 percent and 1 percent significance levels. The regressions also include three seasonal dummies (coefficients not reported). The linear trend takes integer values between 1 and 8 over the 2002:Q1–2003:Q4 period. In columns 2 and 4 the coefficients labeled "constrained" are fixed at their values estimated in column 1.

terminals increases them. The coefficient of the interest rate on checking accounts is negative but not significant. The coefficient of the number of ATM cards is 0.89, not significantly different from 1 (t-statistic of 0.7). The point estimate of the parameter δ is 2.82, statistically greater than 1 at the 5 percent confidence level (t-statistic of 2.2). Thus, the requirement spelled out above is satisfied; the functional form hypothesized in equation (2) finds support in the data.[12]

Altogether, while simple, equation (4) seems to capture some essential features of the demand for ATM withdrawals. Considering that it does not feature a lagged dependent variable on the right-hand side (we experimented with specifications incorporating one, but the related coefficient turned out to be nonsignificant), it tracks the data quite well.

We can now test the null hypothesis—that after the changeover, $p_t = p_t^o$—against the alternative, $p_t > p_t^o$. Since p_t is not observable after 2001:Q4, we consider two hypotheses about its behavior. The first is the following:

$$\begin{aligned} p_t &= p_t^o, && t \leq 2001\text{:Q4}, \\ p_t &= p_t^o + \Delta + \xi_t, && t \geq 2002\text{:Q1}, \end{aligned} \tag{5a}$$

where Δ is a positive constant and ξ_t is a white-noise term independent of η_t, with variance σ_ξ^2. The expressions in (5a) could be an appropriate description of a one-off increase in the true price level after the changeover. An alternative hypothesis is

$$\begin{aligned} p_t &= p_t^o, && t \leq 2001\text{:Q4}, \\ p_t &= p_t^o + gT + \xi_t, && t \geq 2002\text{:Q1}, \end{aligned} \tag{5b}$$

where g is a positive constant and T is a linear trend ($T = 1$ in 2002:Q1, $T = 2$ in 2002:Q2,...). This formulation would entail a

[12]The microdata from the households survey run by the Bank of Italy also suggest that δ is greater than 1. Introducing time subscripts in (2), taking logs of both sides and differentiating, δ can be shown to equal the ratio between the growth rate of consumption expenditure for the average ATM cardholder and the corresponding growth rate for the average consumer. Between 1998 and 2004 this ratio averaged about 1.4, as consumption among cardholders consistently outgrew average per capita consumption at the nationwide level.

widening gap between the observed (official) and the true price deflator, implying a permanent inflation bias. It would be unrealistic for large T but could be appropriate over our sample period, which only covers eight quarters after the changeover. Substituting (5a) into (4) yields

$$\begin{aligned} c_t - p_t^o &= \alpha + \beta d_t^{ATM} + \gamma d_t^Q + \delta(e_t - p_t^o) + \varepsilon r_t \\ &\quad + \lambda n_t + \eta_t, \quad t \leq 2001{:}\mathrm{Q4}, \\ c_t - p_t^o &= \alpha + \theta_0 + \beta d_t^{ATM} + \gamma d_t^Q + \delta(e_t - p_t^o) + \varepsilon r_t \\ &\quad + \lambda n_t + \nu_t, \quad t \geq 2002{:}\mathrm{Q1}, \end{aligned} \tag{6}$$

where $\theta_0 = (1 - \delta)\Delta$ and $\nu_t = \eta_t + (1 - \delta)\xi_t$, with variance $\sigma_\nu^2 = \sigma_\eta^2 + (1-\delta)^2\sigma_\xi^2$. A way to test the null hypothesis of no distortion in the price level after the changeover against the alternative hypothesis (5a) would then entail estimating equation (6) over the entire sample period 1993:Q2–2003:Q4, introducing a dummy variable to allow the constant to change over the last two years, and checking for heteroskedasticity. However, it is easy to check that under the alternative hypothesis (5a), equation (6) is affected by a classic errors-in-variables problem; if $\sigma_\xi^2 > 0$, ordinary-least-squares (OLS) coefficients would be inconsistent and biased toward 0. To circumvent this difficulty, we restrict the parameters in (5a) to take the values estimated over the 1993:Q2–2001:Q4 period, and estimate only the coefficient of the 2002–03 dummy, θ_0, which is an unbiased estimator of $(1-\delta)\Delta$. The results of this exercise are presented in column 2 of table 1. The estimated θ_0 has a value of -0.0062; using this figure and the estimated value for δ (2.82), it can be reckoned that the average inflation rate in 2002 was 0.3 percentage point higher than that computed using the official deflator. However, a one-tail t-test cannot reject the null that θ_0 is 0 against the alternative that it is negative. Column 2 also reports the result of an F-test of the null hypothesis $\sigma_\nu^2 = \sigma_\eta^2$ against the alternative $\sigma_\nu^2 > \sigma_\eta^2$, which should hold based on (5a). Again, the null is not rejected.

As mentioned above, if (5a) were true and the errors-in-variables problem were serious after 2001:Q4, OLS coefficients should be biased toward 0. Therefore, as a further check, we estimate all the parameters of (6) over the entire sample period. The estimated coefficients show that the parameters remain stable (see column 3).

Next, we replicate the exercise for our second alternative hypothesis. Substituting (5b) into (4) yields an equation identical to (6), except that now $\theta_0 = (1-\delta)gT$. Thus, beginning with 2002:Q1, a linear trend with coefficient $g(1-\delta)$ should enter the equation. Also, the error term should display the same form of heteroskedasticity as under hypothesis (5a). Specifications in columns 4 and 5 of table 1 show no evidence consistent with the hypothesis of an increase in the price level after the changeover. In both cases the coefficient of the 2002:Q1–2003:Q4 dummy interacted with the time trend is negative. A calculation based on the larger value, −3.6e–3, together with the estimated value of 2.82 for δ, suggests that the average inflation rate in 2002 was 0.5 percentage point higher than that based on the official deflator. However, even in this case, the coefficient is not statistically different from 0, and the null of homoskedasticity cannot be rejected.[13]

3.3 Exploring the Power of the Statistical Test

Our econometric procedure amounts to a t-test on the coefficient of a dummy in a linear regression. Therefore, the properties of our tests and their statistical power are well grounded in standard asymptotic and small-sample theory. It could be argued, however, that the precision of our estimates is insufficient to generate adequate power, e.g., because of the short sample period. To investigate this hypothesis, we perform a counterfactual exercise; we assume that beginning in 2002:Q1 the true price deflator is higher than the official one. Using (5a), we set $p_t^o = p_t - \Delta$ after the changeover. We then assign numeric values to Δ and reestimate the specifications shown in columns 2 and 3 of table 1. If our tests have sufficient power, the coefficient of the 2002–03 dummy should become negative and significant for relatively small values of Δ. We report the results of this exercise for values of Δ ranging between 0.005 and 0.1, implying that in 2002 inflation was between

[13] While we focused on hypotheses (5a) and (5b) in order to maximize the power of the test, we also tested hypothesis $p_t = p_t^o + \Delta + gT + \xi t$, $t \geq$ 2002:Q1, which nests (5a) and (5b). The estimated coefficients for the 2002–03 dummy and for the time-trend results were not statistically different from 0 (individually as well as jointly).

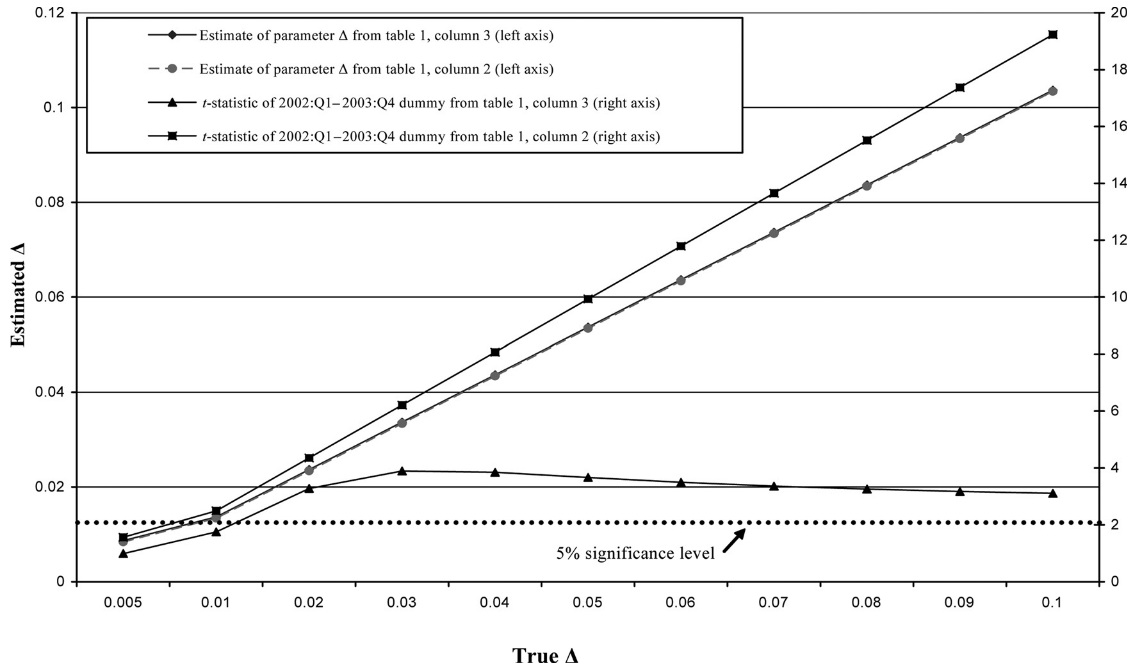
0.5 and 10 percentage points higher than that recorded by official statistics.

Figure 3 plots the "true" Δ, measured on the horizontal axis, against its estimated value, $\hat{\Delta}$, obtained as the ratio between the coefficient of the 2002–03 dummy and $(1 - \delta)$. The curves obtained with the specifications shown in columns 2 and 3 of table 1 virtually overlap, so that they can hardly be distinguished in the figure. They are very close to the 45° lines, indicating that the size of the distortion is captured quite well—in fact, it is systematically slightly overestimated. The figure also shows the precision of the estimates, measured by the t-statistic of the 2002:Q1–2003:Q4 dummy. Both specifications fail to detect the presence of a 0.5-percentage-point distortion. However, the specification in column 2 manages to correctly signal as statistically significant (with a 5 percent confidence level) a value of Δ as little as 1 percent; in this case, the specification in column 3 yields a p-value of .09. Both specifications capture values of Δ greater than 1 percent, at least at the 5 percent confidence level.

3.4 Robustness Check

The above results were subjected to a number of robustness checks. First, we checked the stability of the specification reported in table 1, column 1. The exercises described in section 3.2 entail detecting a structural break in the equation after the fourth quarter of 2001. Thus, it is important that the coefficients in column 1 of table 1 be stable. An obvious candidate for a structural break is the beginning of the single euro-area monetary policy regime, in January 1999. Therefore, the five coefficients α through λ of the specification in column 1 were allowed to change over the 1999:Q1–2001:Q4 period. The F-test of the null hypothesis that the changes in the coefficients are jointly equal to 0 yields $F_{(5,20)} = 2.66$, which does not allow for rejection of the null of parameter stability at the 5 percent confidence level. However, since this value is close to significance, the tests in table 1 were replicated using the equation that allows for parameter change over the 1999:Q1–2001:Q4 period. The results were qualitatively analogous to those in table 1 and therefore are not reported.

Figure 3. The Power of the Statistical Test: Results from a Counterfactual Exercise



Note: The specifications reported in columns 2 and 3 of table 1 were reestimated using a counterfactual price deflator, which jumps by Δ in 2002:Q1. See the text for a detailed description of the exercise. The horizontal axis measures Δ, the shift in the price level that occurred after the changeover, based on hypothesis (5a) in the text. The vertical left axis measures the estimated Δ, computed as $\hat{\Delta}$ = (estimated coefficient of 2002:Q1–2003:Q4 dummy)/$(1-\hat{\delta})$. The t-statistics, measured on the right axis, are reported in absolute value.

Second, the estimates of equation (4) reported in table 1 are subject to a potential endogeneity-bias problem, as some right-hand-side variables (e.g., expenditure) may be simultaneously determined with the dependent variable. Thus, we reestimated the specification in column 1 with two-stage least squares, instrumenting d_t^{ATM}, d_t^Q, $e_t - p_t$, and n_t with their lagged values. The results (not reported) are virtually unchanged.

Third, the analysis in section 3.2 relies on the hypothesis that nominal expenditure e_t is measured without error prior to and after the changeover. However, nominal components of consumption expenditure are computed using both value data (i.e., data measured in nominal terms) and data built from price and quantity indices. Possible mismeasurements in the prices of these components after 2001:Q4 will, in principle, bias e_t as well. Since a detailed breakdown of the data on household consumption by construction method is not available, it is not possible to address this concern in a precise way (see ISTAT 2000). However, ISTAT does publish a breakdown of consumption expenditure used in section 3.2 into two categories: nondurable goods and services. The former is virtually entirely built from value data, and therefore it is not affected by possible mismeasurement in p_t after the changeover. Thus, we rerun the regressions in table 1, proxying e_t with consumption of nondurable goods, i.e., excluding expenditure on services.[14] Table 2, in the appendix, reports the instrumental-variables estimates of this last specification. The elasticity of money demand to real consumption is now 2.3, slightly lower than in table 1 but still significantly different from 1 at the 5 percent confidence level. No appreciable changes in the other results emerge. Similar results were obtained with OLS.

Finally, we rerun our exercises using specifications featuring several alternative combinations of regressors: the number of ATM terminals, the number of POS terminals, the number of ATM cards

[14] Angelini, Ardizzi, and Lippi (2005) also use consumption of nondurables as a proxy for expenditure, in a specification featuring inflation and a time trend among the regressors. The results of the tests are analogous to those reported here.

separately, and/or total households consumption (as opposed to consumption of nondurables and services). In this case as well, the main results of the analysis remain unchanged.

4. Conclusions

Did the euro cash changeover trigger a sudden, substantial increase in the price level in the euro area, largely undetected by the national statistical offices? This paper presented a simple indirect method to address this question for Italy. The basic idea underlying the testing strategy entails searching for the effects that the hypothesized large increase in the price level should have induced on the dynamics of payment instruments—notably, cash withdrawals from ATM terminals. A quarterly data set on aggregate ATM withdrawals and nominal consumption expenditures in Italy was used to test the hypothesis.

The estimation of a demand equation for ATM cash withdrawals, conducted along the lines suggested by a simple theoretical model, allows us to set up econometric tests of the null hypothesis (that after the currency changeover, in the first months of 2002, the official price index continued to measure the general price level correctly), against the alternative (that it underestimated it). The main result of the analysis is that none of the various tests performed provide evidence against the null. Specifically, our point estimates imply that the average inflation rate in 2002 was about one-half percentage point higher than that computed using the official deflator; however, this effect is not statistically significant.

To assess the possibility that failure to reject the null is due to lack of power, we perform a counterfactual exercise: we introduce an artificial increase in the deflator time series beginning in 2002 and rerun our tests. The equation accurately captures the magnitude of the inflation distortion, correctly signaling it as statistically significant as soon as it grows greater than or equal to 1 percent on an annual basis. We conclude that the determinants of the well-documented disconnect between official and perceived measures of inflation cannot be ascribed to a sizable mismeasurement by the national statistical offices of the euro-area countries.

Appendix

Table 2. Alternative Estimates of Equations (4) and (6) ($e_t - p_t$ Measured as Consumption of Nondurable Goods; Instrumental-Variables Estimates)

		Equation (6) Estimated over 1993:Q2–2003:Q4			
		Test of Hypothesis (5a)		Test of Hypothesis (5b)	
	Equation (5) Estimated over 1993:Q2–2001:Q4 1	Estimating Only Dummy Coefficient 2	Estimating All Coefficients 3	Estimating Only Dummy Coefficient 4	Estimating All Coefficients 5
ATM Terminals Diffusion (d_t^{ATM})	2.47**	2.47	2.49**	2.47	2.54**
	4.10	Constrained	6.60	Constrained	4.90
POS Terminals Diffusion (d_t^Q)	−12.49**	−12.49	−12.96**	−12.49	−12.13**
	−5.20	Constrained	−6.60	Constrained	−5.40
Real Consumption (log; $e_t - p_t$)	2.30**	2.30	2.52**	2.30	2.30**
	3.90	Constrained	4.30	Constrained	3.70
Number of ATM Cards (log; n_t)	1.11**	1.11	1.10**	1.11	1.08**
	7.70	Constrained	8.70	Constrained	8.70
Interest Rate on Checking Accounts (log; r_t)	0.01	0.01	0.02	0.01	7.2e–3
	0.40	Constrained	0.80	Constrained	0.30
Constant	−4.40	−4.40	−6.56	−4.40	−3.54
	−0.60	Constrained	−1.10	Constrained	−0.50
Dummy 2002:Q1–2003:Q4	—	−8.5e–3	−3.7e–3	—	—
	—	−0.90	−0.30	—	—
Dummy 2002:Q1–2003:Q4*Linear Trend	—	—	—	−1.8e–3	−2.2e–3
	—	—	—	−1.00	−0.70
F-test for $\sigma_\nu^2 = \sigma_\eta^2$ versus $\sigma_\nu^2 > \sigma_\eta^2$ (p-value)	—	0.89	0.92	0.90	0.96
Number of Observations	34	43	42	43	42
R^2	0.99	—	0.99	—	0.99
DW	1.85	1.69	1.88	1.69	1.88

Note: The dependent variable is the (log of) aggregate cash withdrawals from ATM in real terms. σ_η^2 and σ_ν^2 denote, respectively, the variance of the error term in equation (6) before and after 2001:Q4. Instrumental-variables estimates. d_t^{ATM}, d_t^Q, n_t, and $e_t - p_t$ are instrumented using their own lags. $e_t - p_t$ is proxied by real consumption of nondurable goods. Heteroskedasticity-robust t-statistics are reported below each coefficient. ** denotes 1 percent significance levels. The regressions also include three seasonal dummies (coefficients not reported). The linear trend takes integer values between 1 and 8 over the 2002:Q1–2003:Q4 period. In columns 2 and 4 the coefficients labeled "constrained" are fixed at their values estimated in column 1.

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Appendix. Dataset for Did Prices Really Soar after the Euro Cash Changeover? Evidence from ATM Withdrawals

Date	Consumption by Households				n° Debit Cards (cards)	n° ATM Terminals (tatm)	n° POS Terminals (tpos)
	Non Durable Goods (€ million)		Services (€ million)				
	Current Prices (cgn)	Constant Prices (cgr)	Current Prices (csgn)	Constant Prices (csgr)			
6/30/1993	53580	58810	47423	54141	11200000	15400	64000
9/30/1993	56025	60894	59001	66232	11500000	16000	70000
12/31/1993	59017	63702	47584	52724	11815283	16862	78290
3/31/1994	61353	65316	49097	53479	12111920	17586	86974
6/30/1994	56267	59527	52106	55966	12408557	18309	95657
9/30/1994	58654	61565	65258	69146	12705194	19032	104340
12/31/1994	61905	64308	52631	55048	13001830	19755	113023
3/31/1995	65111	66607	54549	55873	13343690	20289	123499
6/30/1995	60206	60288	57913	58354	13685550	20822	133974
9/30/1995	62953	62578	71596	71076	14027410	21355	144450
12/31/1995	65551	64349	57242	55996	14369270	21888	154925
3/31/1996	68522	66453	58864	56807	14759734	22505	169870
6/30/1996	62853	60125	62801	59831	15150198	23122	184815
9/30/1996	65082	62345	76257	72156	15540662	23739	199760
12/31/1996	66812	63670	60998	57147	15931125	24355	214705
3/31/1997	70264	66609	62462	58052	16242754	24655	229906
6/30/1997	64852	61327	66156	60915	16554382	24955	245107
9/30/1997	67901	64240	79226	72728	16866010	25255	260308
12/31/1997	69842	65702	63891	57994	17177638	25554	275509
3/24/1998	73829	69064	65526	59021	17562023	26176	293020

(continued)

Date	Consumption by Households				n° Debit Cards (cards)	n° ATM Terminals (tatm)	n° POS Terminals (tpos)
	Non Durable Goods (€ million)		Services (€ million)				
	Current Prices (cgn)	Constant Prices (cgr)	Current Prices (csgn)	Constant Prices (csgr)			
6/30/1998	67936	63451	70159	62658	17946408	26798	310531
9/30/1998	70656	65845	84217	74989	18330793	27420	328042
12/31/1998	71531	66549	68244	60057	18715178	28042	345552
3/31/1999	75673	70176	69717	60874	19236422	28614	371059
6/30/1999	70028	64610	74257	64272	19757665	29185	396565
9/30/1999	72505	66416	89187	76925	20278908	29757	422072
12/31/1999	73901	67375	72193	61518	20800151	30328	447578
3/31/2000	77784	70284	74403	62695	20887762	30714	486830
6/30/2000	73478	65921	79843	66803	20975372	31099	526082
9/30/2000	76367	67963	96698	80784	21062983	31485	565334
12/31/2000	77514	68699	77204	63790	21150593	31870	604585
3/31/2001	81327	71720	78759	63959	21575546	32491	640450
6/30/2001	75412	66163	84458	68127	22000499	33111	676315
9/30/2001	77178	67407	102141	82195	22425452	33732	712180
12/31/2001	77811	67854	80934	64288	22850405	34352	748045
3/31/2002	82335	71251	82537	64305	23275393	34875	766116
6/30/2002	76855	66161	88367	68180	23700380	35399	784186
9/30/2002	79830	68232	106954	82283	24125367	35922	802256
12/31/2002	79430	67602	86008	65469	24550354	36445	820326
3/31/2003	84487	71233	86987	65370	24619137	36619	846875
6/30/2003	79487	66725	92857	69473	24687919	36793	873424
9/30/2003	82697	68814	111058	82828	24756702	36967	899973
12/31/2003	82063	67968	89187	66095	24825484	37141	926521

(continued)

Date	ATM Withdrawals		POS Transactions		Interest Rate on Checking Accounts	Quarterly Dummies		
	Number (numatm)	Value (€) (valatm)	Value (€) (valpos)	Number (numpos)	(tadec)	(du2)	(du3)	(du4)
6/30/1993	46381875	6998138684	533253663	5978836	6.94	1	0	0
9/30/1993	49204482	7535682008	524105229	6370769	5.99	0	1	0
12/31/1993	64412127	10056666681	674192197	7572310	5.56	0	0	1
3/31/1994	56151823	8928425261	675794107	7648041	5.35	0	0	0
6/30/1994	65539883	10354520293	775771114	8778908	4.95	1	0	0
9/30/1994	71092061	11488617280	762462081	9354396	4.81	0	1	0
12/31/1994	83495137	13106058548	980806825	11118655	4.97	0	0	1
3/31/1995	73997470	11805478564	932919997	10236389	5.07	0	0	0
6/30/1995	82812000	13130088776	1119560800	11907922	5.42	1	0	0
9/30/1995	83509230	13585357393	1093184842	14768884	5.68	0	1	0
12/31/1995	88836539	13992918115	1491621006	16802025	5.80	0	0	1
3/31/1996	84262001	13589659499	1337167346	15023166	5.76	0	0	0
6/30/1996	90823041	14584011014	1470436451	17005636	5.64	1	0	0
9/30/1996	94036500	15342473408	1484998479	18091024	5.06	0	1	0
12/31/1996	97692229	15785756111	1956987399	23929958	4.60	0	0	1
3/31/1997	93809542	15152752965	1876875638	25389806	4.18	0	0	0
6/30/1997	103469921	16946387649	2201705343	27607953	3.95	1	0	0
9/30/1997	105276524	17109292095	2245122320	30786292	3.48	0	1	0
12/31/1997	106700302	17483278669	3222608396	39667626	3.33	0	0	1
3/24/1998	105567243	17105511111	2619744138	34411957	2.98	0	0	0
6/30/1998	117029204	18731651559	3238467778	44685688	2.60	1	0	0
9/30/1998	117949986	19075922262	3397356760	48378762	2.41	0	1	0

(continued)

Date	ATM Withdrawals Number (numatm)	ATM Withdrawals Value (€) (valatm)	POS Transactions Value (€) (valpos)	POS Transactions Number (numpos)	Interest Rate on Checking Accounts (tadec)	Quarterly Dummies (du2)	Quarterly Dummies (du3)	Quarterly Dummies (du4)
12/31/1998	127465911	20646491964	4155193238	53917354	2.01	0	0	1
3/31/1999	112331696	18053712416	3552739081	50230142	1.35	0	0	0
6/30/1999	123057971	19705867020	4376964764	61294854	1.14	1	0	0
9/30/1999	122802234	19829629694	4027689262	60113776	1.08	0	1	0
12/31/1999	127445989	20622814882	5322057582	75861439	1.15	0	0	1
3/31/2000	120966072	19405002784	4875654509	66775482	1.31	0	0	0
6/30/2000	132552963	21316755325	5667510683	79488985	1.49	1	0	0
9/30/2000	134821254	21847376455	5504233202	78578664	1.77	0	1	0
12/31/2000	136283090	22411152989	7053621011	91673044	2.02	0	0	1
3/31/2001	133985711	21558491184	6519311295	90257739	2.07	0	0	0
6/30/2001	140716944	23210523447	7402217621	102715270	2.02	1	0	0
9/30/2001	142133476	23151055195	7334520562	105452533	1.90	0	1	0
12/31/2001	144832094	22669041391	9319211811	128729304	1.50	0	0	1
3/31/2002	138544369	21891152587	9511826494	119700321	1.36	0	0	0
6/30/2002	147207421	23695856850	9897826582	125713076	1.40	1	0	0
9/30/2002	150569811	24887308290	9897778085	134481649	1.41	0	1	0
12/31/2002	151523226	24251076598	11215147246	145242116	1.31	0	0	1
3/31/2003	142719098	22777218287	9904067728	131609945	1.11	0	0	0
6/30/2003	154296578	24750470871	10573810649	140725556	0.95	1	0	0
9/30/2003	152409746	25025849734	10434503975	144141004	0.76	0	1	0
12/31/2003	149865579	24778483381	11933242131	156144980	0.76	0	0	1

The Great Inflation and Early Disinflation in Japan and Germany*

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This paper considers the Great Inflation of the 1970s in Japan and Germany. From 1975 onward, these countries had low inflation relative to other large economies. Traditionally, this success is attributed to stronger discipline on the part of Japan and Germany's monetary authorities—e.g., more willingness to accept temporary unemployment, or greater determination not to monetize government deficits. I instead attribute the success of these countries from the mid-1970s to their governments' and monetary authorities' acceptance that inflation is a monetary phenomenon. Likewise, their higher inflation in the first half of the 1970s is attributable to the fact that their policymakers over this period embraced non-monetary theories of inflation.

JEL Codes: E52, E58, E64, E65.

1. Introduction

This paper considers the Great Inflation of the 1970s in Japan and the Federal Republic of Germany. These countries are notable for the fact that their peaks in inflation came earlier in the decade—in 1974—than in many other countries, while their inflation rates in

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the late 1970s and early 1980s were well below those prevailing elsewhere. A comprehensive explanation for the policy behavior underlying the Great Inflation needs to account for the experiences of these two countries. I argue that the monetary-policy-neglect hypothesis, previously applied to countries with poorer inflation records, can account for behavior in Japan and Germany too.

It has become a cliché to attribute Germany's success in achieving price stability to greater understanding on the part of German policymakers and the German population of the costs of inflation (with this, of course, frequently said to be the legacy of its hyperinflations), and similar observations were made about Japan after it joined the low-inflation league. But the discussion in this paper will move away from this kind of explanation for success. Germany's policymakers certainly stressed the costs to society of even moderate inflation. But such stress was not something that distinguished Germany from elsewhere; e.g., in the United States, Arthur Burns, Chairman of the Board of Governors of the Federal Reserve System, made many apocalyptic statements about the effects of inflation. Nor is it even clear that Germany's policymakers exhibited a relatively greater willingness to accept the costs of disinflation. Indeed, those who believe that the differences between Germany's inflation performance and that of other countries are attributable to different policymaker preferences may be surprised to find that Helmut Schmidt, Germany's Finance Minister and then Federal Chancellor over much of the 1970s,[1] stated that "we in the Federal Republic can in any case tolerate 5% inflation more than 5% unemployment" (*SZ*, December 21, 1972, *a.t.*).[2] While Schmidt clarified that this claim was "a political value judgment" and did not reflect a belief that such a choice existed,[3] his statement

[1]Despite the Bundesbank's independence, it is appropriate to include the executive branch among the key decisionmakers on monetary policy matters in Germany in the 1970s. The Bundesbank described its 1974 policy changes as "part of a joint strategy agreed with the Federal Government" (*BBAR*, April 1975, p. 1), and the executive branch was the senior partner on key issues such as exchange rate policy.

[2]The abbreviation "*a.t.*" denotes the present author's translation of German material. Appendix 1 provides acronyms for periodicals cited in the text, while appendix 2 gives bibliographic details for specific articles.

[3]Consistent with this, I show below that Germany's policymakers did not subscribe to Phillips-curve trade-off analysis.

puts in perspective the idea that Germany's success emerged from an especially "hawkish" attitude to inflation compared with other countries.

The belief in policy circles that inflation is very costly to the economy and, therefore, to society does not distinguish Germany and Japan in the 1970s from countries like the United States and the United Kingdom. Rather, the resistance to nonmonetary views of inflation is what makes these countries unusual. As we will see, these views had only a brief heyday in each country (roughly 1971–72 in both cases) and even then, there was successful resistance to the idea of imposing national wage and price controls.

This distinguishing feature is consistent with the monetary-policy-neglect hypothesis. The message of that hypothesis is that high inflation is the outcome of episodes during which policymakers attributed inflation to nonmonetary factors, delegating inflation control to nonmonetary devices (such as wage and price controls). Disinflations and low-inflation periods typically follow policymakers' acceptance of the monetary view of inflation and their resulting adjustment of monetary policy. By extensive analysis of statements by policymakers and of key economic commentary during the 1970s, I show that Japan and Germany's inflation-disinflation pattern fits this story.

The experiences of Germany and Japan are natural to study jointly. In the pre-inflation-targeting era, Japan's monetary policy, alongside Germany's, was regarded as an international benchmark. This perspective on Japan's record was reflected in the title of a 1981 paper by a senior Bank of Japan official, "Why Is the Performance of the Japanese Economy So Much Better?,"[4] and Taylor's (1993, 5) observation that "Japanese monetary policy did deliver a low inflation rate much earlier than the other countries, and apparently was doing something 'right.'" Over this period, Japan was Milton Friedman's favorite example of successful monetary policy (see, e.g., Friedman 1983, 1990; and *NW*, September 4, 1978). I argue that the key factor behind this achievement was that Japan's emphasis on nonmonetary means of fighting inflation was brief and over by 1973.

[4]Suzuki (1981).

My emphasis on the doctrines guiding policy, and my specific focus on the 1970s, complements the study of U.S. policy by Romer and Romer (2002) and also distinguishes my study from most existing work on Germany and Japan. While there are many studies of Bundesbank monetary policy, relatively few contributions cover the major years of Germany's period of inflation. Clarida and Gertler (1997) only briefly consider the period before 1978; indeed, the earliest statement by the Bundesbank that they consider is from 1989. Bernanke and Mihov (1997, 1026) explicitly limit their coverage to that following "the inception of the current regime," dated as the beginning of 1975. Similarly, Issing (2005) focuses mainly on post-1977 policy developments, which accounts for his characterization of Germany as not having experienced the Great Inflation.

Von Hagen (1999a) does cover the leadup to monetary targeting, but his discussion concentrates on the conflicts between fiscal and monetary policy, and does not consider the debate over monetary versus nonmonetary views of inflation, which had largely been resolved by the time monetary targeting was adopted. By contrast, I study Germany's Great Inflation and its disinflation, much of which preceded 1975, with emphasis on the conflict between monetary and nonmonetary perspectives on inflation control. This emphasis brings out an element of Germany's experience highlighted by the Bundesbank in 1980 when it noted that high inflation had arisen during a period of "underestimation of monetary policy as an economic policy instrument."[5]

Turning to Japan, there are antecedents to my analysis in the form of the aforementioned Friedman discussions, as well as Hetzel (1999). Friedman (in *NW*, September 4, 1978) discusses Japan's "fundamental change in monetary policy" in 1973, which amounted to an acceptance that "[s]ubstantial inflation is a monetary phenomenon." Hetzel studies Japan's postwar monetary policy and makes the important observation that Japan's 1970s disinflation produced a "profound change in professional and popular views" (1999, 7), discarding nonmonetary views of inflation, a conclusion supported by the analysis provided here. But neither Friedman nor Hetzel provides specific documentation of these changes in views. The discussion in

[5]Deutsche Bundesbank (1980, 291).

this paper fills this gap by drawing on coverage of inflation in Japan in several newspapers during the 1970s.

In addition, while both Friedman and Hetzel emphasize that a floating exchange rate enabled Japan to disinflate, they do not discuss why Japan chose disinflation when other countries, such as the United Kingdom, initially chose monetary expansion after floating their exchange rate. I provide an answer by studying developments in Japanese macroeconomic debates. Furthermore, the only policymaking agency Hetzel discusses is the Bank of Japan. This is problematic in studying the 1970s, because (i) the Bank of Japan was not independent over this period, so senior members of the Japanese government were key makers of monetary policy,[6] and (ii) when nonmonetary views of inflation guide policymakers, some of the major policy mistakes will take the form of attempting nonmonetary strategies against inflation. My focus on a wider range of policymakers and policy agencies overcomes this limitation.

The analysis here also sheds light on the merits of accounts of Japan and Germany's success that emphasize nonmonetary factors. There is wide acceptance among monetary economists that differences in monetary policy account for different countries' inflation experiences during the 1970s. Nevertheless, adherents to nonmonetary views of inflation have offered their own rationalizations for the price stability observed in Germany and Japan. For example, in 1977 Denis Healey, the UK Chancellor of the Exchequer, said, "If you talk to West German Chancellor Helmut Schmidt about his country's successes, he will say that the moderation of the unions in Germany, in limiting their wage demands, is largely due to the political relationship established between the Government and the unions" (*SUN*, March 9, 1977). Walter Heller, former chairman of the Council of Economic Advisers, testified to a U.S. Senate committee in 1979 that Germany's low inflation reflected the fact that Germans were "benefiting from what they call their 'Concerted Action,' from a kind of social contract or compact between business and government and labor."[7] Similarly, Braun (1986, 240) claims that the

[6]Friedman (*NW*, September 4, 1978) acknowledges the Finance Minister's role in producing the 1973 change in monetary policy.

[7]Walter Heller, in his March 5, 1979, testimony to the U.S. Senate Budget Committee (1979, 47).

"Concerted Action" incomes policy in Germany "proved to be useful in promoting wage moderation in 1973–75," while a former U.S. ambassador to Germany asserted that moderate union behavior in German wage negotiations "lowered cost-push [pressures] and was certainly an important reason for the relatively favorable inflation rate."[8] Discussions of "Japan Inc." during the 1970s adopted an analogous line of argument for Japan, and in this spirit a Tokyo economics columnist observed in 1978, "Much of Japan's success in fighting rampant inflation has been ascribed ... particularly to the Japanese version of an 'incomes policy' designed to restrain wage increases" (*JT*, October 16, 1978). I provide evidence on whether these frequently cited features of German and Japanese economic policy made any material contribution to fighting inflation.

This paper proceeds as follows. Section 2 discusses the monetary-policy-neglect hypothesis. Section 3 discusses the methodology that I use to study policymaking in Japan and Germany. Section 4 covers Japan, and section 5 turns to Germany. Section 6 relates the two countries' experiences to "trade-off exploitation" explanations for the Great Inflation. Section 7 concludes.

2. The Monetary-Policy-Neglect Hypothesis

Consider an expectational Phillips curve in generic form:

$$\pi_t = \pi^e + \alpha(y_t - y_t^*) + u_t, \tag{1}$$

where π_t is quarterly inflation, π^e is expected inflation, $y_t - y_t^*$ is the output gap, and u_t is a shift factor (a cost-push shock). Written in terms of equation (1), the Phillips curve delivers special cases such as a traditional shift-adjusted expectational Phillips curve (as in, e.g., Humphrey 1985), where π^e is $\mathrm{E}_{t-1}\pi_t$, or the New Keynesian Phillips curve augmented by a cost-push shock—used by Clarida, Galí, and Gertler (1999)—where π^e corresponds to expected future inflation $\mathrm{E}_t\pi_{t+1}$. The latter version of the Phillips curve allows the u_t shock to be serially correlated.[9] It is clear that if u_t is serially correlated, it matters for both inflation and expected future inflation,

[8] Hillenbrand (1983, 25, 27).

[9] In the traditional Phillips curve corresponding to $\pi^e = \mathrm{E}_{t-1}\pi_t$, a serially correlated u_t cannot be contemplated in general.

and becomes the sole determinant of these two series if $\alpha = 0$. If we consider the New Keynesian Phillips curve further and generalize to allow for a constant term, as well as the usual $\beta < 1$ coefficient on expected inflation, it can be shown that the expression for expected future inflation is

$$\mathrm{E}_t\pi_{t+1} = K + \alpha \mathrm{E}_t \Sigma_{i=0}^{\infty}\beta^i \left(y_{t+i+1} - y^*_{t+i+1}\right) + (1-\beta\rho_u)^{-1}\rho_u \hat{u}_t, \quad (2)$$

where K is a constant, ρ_u is the AR(1) coefficient for the exogenous u_t series, and $\hat{u}_t$ is the deviation of u_t from its mean.

Nonmonetary and monetary views of inflation deliver rival sets of restrictions on equation (2). The *monetary* view asserts that $\rho_u = 0$ (implying that $\hat{u}_t$ does not matter for expected future inflation), that $\mathrm{E}[u_{t+k}] = 0$ for all k (and so expectations of u do not matter for the constant term K), and that $\alpha > 0$ whatever the value taken by the output gap. The monetary view of inflation thus attributes the 1970s inflation to excess demand and gives cost-push shocks no role other than as one-time price-level shocks (which, for a given expected path of the output gap, matter for current inflation but not expected future inflation).

The *nonmonetary* view of inflation, by contrast, contends that $\mathrm{E}[u_{t+k}]$ is generally nonzero and that high inflation reflects high current and prospective values of u_t; that $\rho_u > 0$; and that $\alpha = 0$ when the output-gap sum in equation (2) is negative. The nonmonetary view of inflation thus attributes the 1970s inflation to cost-push shocks and implies that creating negative output gaps does not remove inflationary pressure.

The monetary-policy-neglect hypothesis states that the monetary view of inflation is the correct one and that high-inflation episodes during the 1970s were the result of policymakers' embrace of the nonmonetary view of inflation. This hypothesis has previously been applied to countries whose inflation rates were generally high throughout the 1970s,[10] but it has implications for low-inflation experiences too. According to this hypothesis, countries that experienced relatively low inflation, such as Japan and Germany from 1975 onward, did so because their policymakers converted early to a monetary view of inflation. The remainder of this paper documents the

[10]Nelson (2005) provides evidence supporting the hypothesis for the United States and the United Kingdom.

case for the monetary-policy-neglect hypothesis as a description of these two countries' experiences, as well as pointing out weaknesses of alternative hypotheses.

3. Methodology

My procedure in this paper is to draw on public statements by key policymakers, using these statements to deduce their implied model of inflation. Newspaper reports are used as a major source for policymaker statements. This methodology raises two major questions: (i) How *reliable* are the public statements as an indicator of policymakers' true views about the economy? (ii) How *representative* are the statements that I present—i.e., are my findings insulated from selection bias? I consider each of these issues in turn.

How reliable are the statements? An objection that could be raised about my reliance on public statements for deducing policymaker beliefs is that policymakers give different views publicly from those they express privately. This objection is, however, unlikely to be valid for the type of policymaker views I consider, which pertain to how the structure of the economy behaves over periods of a quarter or more. Policymakers have no plausible incentive to be secretive about matters like this. They certainly may not be forthcoming on specific day-to-day considerations about policy tactics, such as the timing of forthcoming interest-rate decisions. But strategic thinking, reflecting policymakers' longer-term macroeconomic judgments about how the economy works, is a matter about which heavy disclosure is likely. Policymakers, both in the 1970s and today, want the public to know their thoughts about the causes and costs of inflation, and about the links between economic management and economic outcomes.[11]

How representative are the statements? There are several reasons for being confident that the material I present gives a representative picture of policymakers' views. First, my procedure of

[11]This contention is supported by the contents of material on 1970s economic policy in the United States that has been declassified since my 2005 article. That article applied to U.S. policy the same methodology that I use in this paper. The 1978 Federal Open Market Committee transcripts, released in 2007, support my characterization of Federal Reserve Chairman G. William Miller's views based on his public statements.

looking at contemporaneous statements by policymakers automatically avoids the risk of relying on ex post rationalizations that might appear in retrospective accounts by former policymakers. The potential unreliability of retrospective accounts is illustrated by the case of Federal Reserve Chairman Arthur Burns in the United States. Romer (2005, 181) judges that Burns's 1979 account of his 1970–78 period as Chairman contains "a substantial amount of wishful revisionism." In addition, it is not widely known that Burns (1978), often cited as a complete collection of Burns's public statements as Chairman, is actually a very partial collection, omitting some of the most unorthodox remarks about inflation that Burns made over 1970–78 (including an item from 1977, quoted in section 5 below). The problem of revisionism extends, of course, beyond the specific example of Burns. My methodology overcomes this problem by using contemporaneous material instead of long-after-the-fact accounts.

Second, by relying heavily on newspaper reports, I in effect pool multiple sources of information about policymakers' views. Consider the case of Japan. Bank of Japan statements are useful for providing technical details about the thinking behind policy choices; but Bank of Japan publications, besides often not being available in English, may not give adequate coverage of statements by members of Cabinet, who were the most senior policymakers. But newspaper accounts provide coverage of the statements of many policy figures, including both Cabinet members and Bank of Japan officials.

Third, while no account of 1970s developments can hope to provide a completely exhaustive collection of policymaker statements about inflation, the hypothesis that I advance is not one that lends itself to selection bias. The reason is that if, as I argue, German and Japanese policymakers initially subscribed to a cost-push view of inflation, this implies that they *could not* have been guided by the sorts of views prominent in other explanations for 1970s behavior. Embracing a cost-push view of inflation means that one does not believe in a link between the output gap and inflation. Therefore, if policymakers subscribed to a cost-push view of inflation, the following are ruled out as descriptions of policymaker behavior: that policymakers attempted to exploit an inflation/unemployment trade-off, that the absence of a disinflation reflected sacrifice-ratio calculations on the part of policymakers, or that the monetary authorities deliberately permitted inflation by consciously accommodating nonmonetary shocks.[12] The monetary-policy-neglect hypothesis therefore cannot be lumped in with most other hypotheses as part of a "portmanteau" explanation of 1970s policies. If you believe in the monetary-policy-neglect hypothesis, you cannot endorse most of the other hypotheses, even as partial explanations. In line with this contention, I show below that policymakers' embrace of cost-push views was frequently accompanied by their rejection of other views of inflation behavior.

Finally, in the material from which I have obtained the quotations used in this paper, there exist many alternative quotations, carrying the same message, which could be substituted for the ones I present. Selection of material for this paper does not in practice mean excluding information that contradicts my hypothesis; on the contrary, space limitations confine me to presenting only a subset of the material that supports my hypothesis.

All in all, there are grounds for considerable confidence that the statements quoted here are representative of official views on inflation during the 1970s and would not be overturned by a more exhaustive presentation of policymaker statements.

4. Japan

This section studies Japan's Great Inflation and disinflation in detail. The documentary source used to obtain contemporaneous statements by Japan's policymakers is principally the Tokyo English-language daily newspaper, the *Japan Times*, which during the 1970s also provided translated excerpts from other Japanese dailies. In addition, I draw on coverage of Japanese economic policy that appeared in the Asia-region newspapers *South China Morning Post* (Hong Kong) and *Straits Times* (Singapore), as well as material

[12] The implication runs in the reverse direction too. For example, a policymaker who believes that higher inflation is the price that must be paid to buy lower unemployment is in effect subscribing to the belief that inflation is sensitive to the output gap, and so is rejecting the cost-push view of inflation.

in newspapers from the United States, the United Kingdom, and other countries.

4.1 1969–73: Increasing Monetary Policy Neglect

At the end of the 1960s, Japan remained on a completely fixed exchange rate. What domestic policymakers thought about how to control inflation nevertheless mattered greatly—first, because there was considerable monetary policy autonomy in practice despite the fixed exchange rate; second, because erroneous views about inflation behavior meant that the implications of the fixed-exchange-rate regime for inflation control were misunderstood.

In this light, it is significant that as of late 1969, Japanese policymakers characterized the control of inflation as largely separate from monetary policy. Prime Minister Sato said in the Diet: "The stabilization of consumers' prices is the important task for protecting the national livelihood, and it is here that the Government has devoted its greatest effort. While restraining as far as possible the prices of public utilities, I intend to stabilize the consumers' prices through strong policy drives for further growth in productivity, mobility of the labor force, and liberalization of imports" (*JT*, December 2, 1969).

While denying that excess demand currently existed, the authorities did acknowledge the prospect of an excess emerging and thereby *becoming* a source of inflationary pressure. The Bank of Japan's discount rate was raised to 6.25 percent in September 1969, with Finance Minister Fukuda citing "the pace of demand expansion" and the risk of overheating as the reason for the change (*AUP*, October 7, 1969). While it may seem jarring to see domestic factors alone given as the reason for the interest rate increase, it is true that foreign-exchange controls gave Japan's authorities considerable liberty in manipulating domestic interest rates while maintaining a fixed exchange rate.[13] Apparently, however, policymakers were satisfied that this single tightening was sufficient; after the 1969 increase, the discount rate was held constant until October 1970, when it was

[13]Consistent with this, Rasche (1990, 35) observes that there are very large discrepancies in the behavior of short-term market interest rates across the United States and Japan over the quarter-century 1956–80, a period that includes the 1969–80 period studied here.

reduced to 6 percent. The reason for this reduction, it was reported, was that Japan's policymakers believed that monetary restriction had achieved its purpose of slowing down the economy (*JT*, October 28, 1970).

The *Japan Times* editorialized in February 1970 that wage increases in the preceding four years had been "determined by the strong-arm tactics of labor unions.... If this trend continues ... [it] may create a serious 'cost inflation'" (*JT*, February 1, 1970). The government was likewise disposed to analyzing inflation in terms of unit-cost developments but at this stage was less inclined to appeal to wage push as a source of inflation. Vice Minister Kashiwagi expressed a relaxed view: "I anticipate no difficulty because of this rate of wage increase, for worker productivity will increase [by] up to 15 per cent a year and industry therefore will be able to offset the wage increase" (*AZR*, May 13, 1970).

In July 1970 Miyohei Shinohara, an official of the Economic Planning Agency (EPA), called for an incomes policy to cover both wage and nonwage incomes (*JT*, July 10, 1970), and in December the EPA cited wage push as a source of prospective stagflation in Japan. This was noted as "the first time that the danger of a 'cost-push' inflation has been warned in an official government document"(*JT*, December 5, 1970). It would be inaccurate, however, to say that cost-push views had not guided official policy by this point; the government's efforts in 1969 to restrict increases in public services' prices were informed by cost-push analysis, and this approach continued in December 1970 with an indefinite freeze on public charges (*JT*, December 10, 1970). The elevation of wage push to the top of the government's list of cost-push factors was confirmed when Prime Minister Sato himself cited wage push. He signaled that a formal incomes policy was an option: "I am afraid the Government might have to adopt an incomes policy under the circumstances.... An incomes policy never has succeeded anywhere in the world, but as prices will not become stabilized as long as large pay raises continue, I would like to work out some countermeasures" (*SCMP*, December 12, 1970).

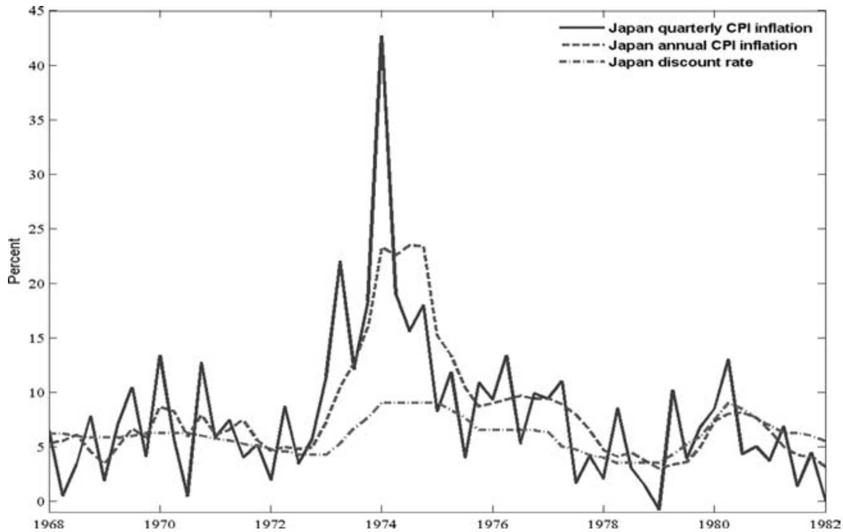
From a monetary perspective on inflation, the really urgent countermeasure Japan needed was greater exchange rate flexibility, a precondition for an assured monetary policy tightening within Japan. But Bank of Japan Governor Sasaki denied that revaluation would help, claiming that the German experience confirmed this (*JT*, March 11, 1971). The opposition to revaluation was rendered moot by what in Japan was labeled the "Nixon shock": the measures that included U.S. dollar devaluations in August and December 1971. The Bank of Japan cut interest rates over this period, with the discount rate in early 1972 standing at 4.75 percent, the lowest level since 1948 (*JT*, May 31, 1972). As well as being aimed at restraining the exchange rate, these cuts had a domestic motivation: the *Japan Times* noted that the government was attempting "to take up the slack in the private sector of the economy ... [via] a low rate [of] interest policy" (*JT*, January 4, 1972). The *Times* claimed that any inflationary impact of such stimulus would be precluded by the "big excess capacity in the economy" (*JT*, January 14, 1972); indeed, it said that this policy might *help* inflation by cutting business costs. In late 1972, Governor Sasaki expressed satisfaction that the economy was recovering but not overheating (*JT*, December 25, 1972), and it was not until February 1973 that a Bank of Japan official said that the output gap was now almost closed (*JT*, February 8, 1973).[14]

4.2 1973–74: Monetary Tightening

Two events combined to create the conditions for a significant monetary policy tightening from March 1973: first, the collapse of the remaining Bretton Woods arrangements, and second, the recognition that the economy was overheating. Even according to the nonmonetary view of inflation, a positive output gap produces inflationary pressure and justifies a tightening of aggregate demand. As noted above, under the fixed-exchange-rate regime, Japan's policymakers retained some discretion with respect to domestic interest rates. But over 1971–73 there had been repeated cuts in these rates, so monetary policy had, if anything, reinforced the tendency for the fixed

[14]Ueda (1993, 193) confirms that the Bank of Japan relied on output-gap measures that underestimated the strength of demand in 1972–73. Since the Bank in early 1973 realized that the gap was closing rapidly, it at least had a more accurate estimate of the output gap than those provided at the time by the Organization for Economic Co-Operation and Development (OECD) (1973, 22), which gave Japan's output gap as −5.9 percent for 1972 on average and as −5.0 percent for the second half of 1972.

Figure 1. Inflation and the Discount Rate in Japan

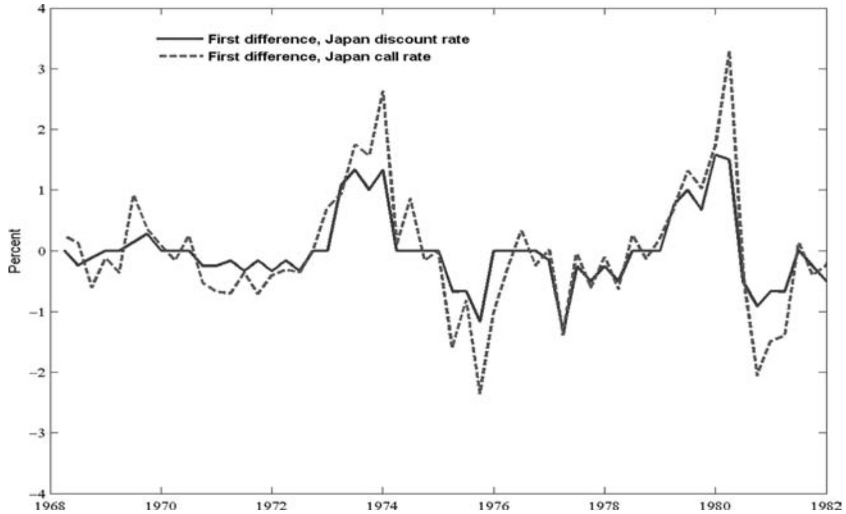


Note: Calculated from Haver–*International Financial Statistics* data.

exchange rate to generate monetary expansion. This ease now began to be removed. The Bank of Japan's discount rate was increased by 75 basis points in March 1973, and Prime Minister Tanaka approved another increase in May 1973 with the statement that he would "tighten the credit squeeze in order to restrain total demand" (*JT*, May 26, 1973). As one Japanese news commentary put it: "Now that the prerequisite [of fixed exchange rates] is gone, the central bank focuses its attention on curbing inflation" (*NKS*, April 11, 1973).

Though a record in nominal terms, the discount rate was well below inflation throughout 1973 and subsequently (see figure 1). But more important was the direction of monetary policy implied by discount-rate choices. Not only did market rates such as the call rate have a considerable spread above the discount rate, but *changes* in the discount rate seem to have had a more than one-for-one impact on the call rate during the 1973–74 monetary tightening (figure 2). Combined with their impact on other asset prices such as the exchange rate, the 1973 discount rate increases packed

Figure 2. Short-Term Interest Rate Changes in Japan



a considerable punch, confirmed also by the sharp slowdown in M2 growth from 1973.[15]

The Japanese government was, however, by no means convinced that eliminating demand pressure was all that was needed to curb inflation, and in April 1973 the government announced several measures inspired by the cost-push view, including increased surveillance of prices of both domestic and imported goods (*JT*, April 14, 1973). Supporting this eclectic approach, the president of Tokai Bank was reported as favoring "nonmonetary measures in order to curb inflation." Such views, the *Japan Times* reported, were in line with "the belief held by many economists here that the Japanese economy is faced with 'composite inflation' "—i.e., cost-push alongside the excess-demand problem—"which cannot be suppressed by conventional monetary instruments" (*JT*, May 30, 1973). Monetary tightening was seen as helpful in moving the output

[15]The fact that other key yields moved in the same direction as the discount rate over the 1970s, albeit not proportionally, supports the focus on that rate in the discussion here. That is, while many asset prices matter for aggregate demand, their reaction to monetary policy operations is assumed to be approximately collinear with that of the short rate. This approximation is also that taken by West (1993) in studying the post-1973 period.

gap to zero and thereby removing excess demand as a contributor to inflation; but, it was believed, this would still leave inflation in the system that had to be attacked with nonmonetary instruments. Always implicit in this position was the view that negative output gaps did not exert negative pressure on inflation. The *Nihon Keizai Shimbun* newspaper[16] endorsed this view, claiming that "it is no longer possible to curb inflation by tight money alone" (*NKS*, May 30, 1973).

In July 1973, Prime Minister Tanaka said he expected the monetary tightening to bring inflation back under control during October–December, and if this did not occur, he would consider a wage freeze (*JT*, July 6, 1973). His overconfident statement reflects lack of appreciation for the length of lags between monetary policy actions and inflation. What is more, he was making a statement that, if proved wrong, would reduce public confidence in monetary tightening as the solution to inflation.

This unwise statement did not, however, lead to a change in monetary policy strategy. When, in November, high inflation was still proceeding, Tanaka judged that demand remained too strong (*JT*, November 2, 1973). In December 1973, he said categorically, "The present situation does not warrant the adoption of an incomes policy" (*JT*, December 9, 1973). The government's reaction in December 1973 to continuing inflation and the OPEC oil shock did include some nonmonetary attempts to fight inflation, such as a freeze on rice prices and rail fares (*JT*, December 22, 1973*a*). But a prominent role was given to what the government called "utmost efforts ... to curb total demand" (*JT*, December 22, 1973*a*), including a 2 percent increase in the discount rate to a record 9 percent (*JT*, December 22, 1973*b*), implying also a sharp increase in the nominal call-money rate (see figure 2).

4.3 1974–75: On the Brink of a U-turn

While following restrictive demand policies, the government continued to hold the position that wage push could produce an independent effect on inflation. Prior to the 1974 wage negotiations,

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[16]According to the *New York Times*, the *Nihon Keizai Shimbun* was "Japan's leading economic newspaper" during this period (*NYT*, June 12, 1974).

Prime Minister Tanaka urged labor to "exercise restraint," stating, "Excessive wage increases . . . invite higher prices" (*JT*, February 22, 1974). *The Economist* claimed that it was "hard to see where prices will stop" unless an incomes policy was introduced (*TE*, January 5, 1974), and when the outcome of the spring wage negotiations was 30 percent nominal-wage growth, an economist at the Japan Development Bank said that Japan was now approaching "fast-paced cost-push inflation" that aggregate-demand measures could not cure (*JT*, June 21, 1974). Similarly, Finance Minister Ohira said in August 1974 that cost-push pressure, manifested in the "vicious circle of prices and wages," was Japan's main economic problem (*JT*, August 3, 1974).

In January 1974 Japan's annual CPI inflation passed 20 percent. It stayed above 20 percent throughout the year. In response, cost-push views strengthened in policy circles and threatened to produce a major policy change in late 1974. In October 1974, the Cabinet set a goal of limiting inflation to 15 percent, with the Prime Minister's Office seeing restraint on public utility prices as a key weapon (*JT*, October 12, 1974). These had featured in anti-inflation packages in previous years; their continued use did not in themselves augur a major policy change. But Prime Minister Tanaka also indicated he would not rule out introducing wage controls, and he was quoted as wishing "to sever the vicious cycle of prices and wages, and promote harmony between the wage problem and the whole national economy" (*SCMP*, October 8, 1974). More promisingly, these official pronouncements were not at the expense of continuing monetary restriction; rather, as the *Japan Times* put it, there remained "a fairly solid and welcome consensus among Government leaders that the restrictive demand management policy should be continued for the supreme objective of slowing inflation" (*JT*, October 23, 1974). Later in October, however, the Bank of Japan's governor said, "We have reached a point where money can't be tightened any further" (*JT*, October 29, 1974), and in December Deputy Prime Minister Fukuda said that Japan's inflation was now "cost-push" (*JT*, December 19, 1974).

The decline in inflation in 1975 (see figure 1) was sufficiently dramatic to hold off a shift to a nonmonetary strategy against inflation, and official statements during 1975 took a more orthodox tone. In particular, demand restriction was reaffirmed as effective against inflation, and a negative output gap was seen as removing inflationary pressure. Deputy Prime Minister Fukuda said that while production was stagnant, "the problems of prices and wages are still a matter of great concern and, therefore, the demand-curbing policy should be continued" (*JT*, January 18, 1975). Similarly, Prime Minister Miki said that while "there is no denying the fact that business is becoming increasingly stagnant," continuing inflation meant "it is not feasible to lift the current restrictive measures on total demand" (*JT*, January 25, 1975). While the discount rate had not been raised during the very large increase in inflation in 1974, neither was it reduced during the first quarter of 1975, by which time statistics were indicating that the economy had contracted during 1974 and that annual inflation was falling rapidly. The fall in inflation was generally recognized as a reaction to the demand restraint: e.g., a correspondent for Singapore's *Straits Times* observed, "Japan's determined anti-inflation campaign is beginning to slow the rise in prices. . . . [T]he inflation rate . . . seems to be coming down because of the recession" (*STR*, January 29, 1975). This was a breakthrough because the recession was widely seen, including in official estimates,[17] as having turned the output gap sharply negative—a "deep slump" in the *New York Times*' estimation (*NYT*, June 12, 1974). Therefore, to attribute the inflation decline to weak aggregate demand was to reject the cost-push position that negative gaps do not pull down inflation.

4.4 1975–79: Entrenching the Monetary Control of Inflation

The first change to Japan's discount rate since 1973, a 0.5 percent cut to 8.5 percent, was made in April 1975. Bank of Japan Governor Morinaga announced the cut with the following caveat: "The discount rate cut does not mean a drastic policy shift; only the signal has changed, from red to reddish yellow" (*NKS*, April 16, 1975). This was a modest cut compared with that being urged by the Prime Minister's Price Stabilization Council, which had advocated a 150-basis-point reduction and had argued that such a move would be anti-inflationary by cutting business costs (*JT*, April 21, 1975). The

[17]One such official estimate was in an Economic Planning Agency report released in August 1975 (*STR*, August 19, 1975).

maintenance of monetary restraint reflected the fact that cost-push views like this were continuing to lose influence. But interest-cost-push views did creep into Deputy Prime Minister Fukuda's analysis when he justified an additional 50-basis-point cut in June 1975 as one that would stimulate the economy and reduce pressure on prices (*JT*, June 9, 1975).

The predominant trend continued to be in the direction of further endorsement of the monetary view of inflation. The president of the Bank of Tokyo said in May 1975 that the "policy of restraining aggregate demand, especially on the monetary front, that had been pursued since the beginning of 1973, has resulted in a pronounced slowing of price advances this year" (*JT*, May 31, 1975). The government white paper on the economy attributed 1973–74's inflation largely to excessive demand and said that inflation had been reduced at the cost of recession (*JT*, August 9, 1975; *JT*, August 14, 1975). Further evidence of the impact of the monetary restriction came in late November, when it was revealed that nationwide annual CPI inflation had fallen below 10 percent in October, the first single-digit outcome since early 1973 (*JT*, November 29, 1975).

By March 1976, it was clear that the government had achieved its goal of bringing down inflation to single digits for the 1975–76 fiscal year (*JT*, March 29, 1976), and its aim for 1976–77 was to bring inflation down further to a maximum of 8.6 percent (*JT*, January 29, 1977). From a monetary perspective on inflation, a decline of this magnitude was close to being locked in by the prior period of monetary restraint and subsequent permanent reduction in money growth. From a cost-push perspective, there was no such guarantee, and the *Japan Times*, taking this approach, was pessimistic: "Cost-push inflationary pressures remain strong.... There is indeed a possibility that we might see the inflation rate soar to a double-digit level again" (*JT*, November 15, 1976). Inflation for the fiscal year ending in March 1977 ultimately came in at 9.4 percent (*JT*, April 29, 1977), an interruption of the decline since 1974. The government set a target of 7 percent for fiscal year 1977–78 (*JT*, April 28, 1977). The overshoot of the 1976–77 target proved to be an aberration, out of line with the monetary restraint observed since 1973; it was compensated for by the rapid decline in inflation in 1978. In July 1978, the *Japan Times* observed, "Perhaps the most striking development on Japan's domestic economic scene today is regained price stability" (*JT*, July 26, 1978), with CPI inflation around 4 percent, well below the 7 percent maximum target. Nevertheless, the article acknowledging this achievement attributed it to the recent yen appreciation rather than the turnaround since 1973 in monetary policy.

All these developments were against the background of the major post-1973 productivity slowdown. Much commentary in Japan in 1973–74 accurately saw the economy as undergoing a permanent shift in its trend growth rate. For example, the *Japan Times* editorialized at the end of 1973 that it was "broadly discerned that the country must save energy consumption and settle for a much lower rate of growth in the future" (*JT*, December 30, 1973), while Finance Minister Fukuda said in May 1974 that in coming years Japan would seek a growth rate "acceptable by international standards" and could not return to the 15 percent growth rates experienced in the past (*JT*, May 10, 1974). So estimates of the output gap in the years following 1973 were not on the grossly erroneous scale that would have resulted from assuming no change in the trend of potential output in years after 1973,[18] but the magnitude of Japan's slowdown did mean substantial output-gap errors. An OECD report released in mid-1977 gave Japan's output gap as about −13.5 percent as of the end of 1976 (McCracken et al. 1977, 84). This was likely more pessimistic than the Japanese government's estimates of the output gap, for the OECD, unlike the government, did not acknowledge any quarter of positive output gaps in Japan since 1970, not even in 1973.[19] Within Japan, the *Japan Times* in late 1977 said there was a "huge surplus of productive capacity" of "well over 10 trillion yen" (*JT*, September 8, 1977)—i.e., an output gap of at least −6 percent. The government itself set a real-GDP-growth target of 7 percent for 1978–79 (*JT*, January 22, 1978) and, in pursuit of

[18]Thus official estimates were not as severely in error as those later given by Brown and Darby (1985, 71–75), who estimated potential output by fitting a log-linear trend to Japan's real GDP over 1952–79. Because it was heavily affected by the pre-1974 trend, Brown and Darby's procedure resulted in a negative and continuously worsening output-gap series for Japan over 1976–79, with the 1979 output gap more negative than −12 percent. In contrast to this series, estimates of the output gap used by Japanese authorities in the late 1970s did record that the output gap was becoming less negative over 1978–79.

[19]See Laidler (1978, 1043).

this, the discount rate continued to be cut, to 3.5 percent in early 1978, with implied bank lending rates of about 5.9 percent (*JT*, March 16, 1978; *JT*, June 23, 1978). In October 1978 the administrative councilor of the Economic Planning Agency gave Japan's output gap as −11 percent as of the previous March (*JT*, October 6, 1978).

Despite this perceived large gap, the authorities in April 1979 increased the discount rate for the first time since 1973. Analyzing this increase, the *Japan Times* said that "the Bank of Japan let it be known on Monday that it is determined to fight a resurgence of inflation" (*JT*, April 18, 1979), with the authorities in particular wishing to avoid excessive money growth of the kind observed in 1972–73, and to forestall a continuation of rising wholesale price inflation. Discussing the increase in wholesale inflation, an official government bulletin cited rising oil and commodity prices but also acknowledged that "improvements in the supply and demand situation"—that is, a diminishing output gap—were a contributing factor (*JT*, June 26, 1979). This contrasts with the authorities' position in 1972, when they did not regard a narrowing of the gap as a signal for tightening. The change in the gap was now given weight in policy decisions—thus embedding into monetary policy a "speed-limit" dimension, in the terminology of Walsh (2003).[20]

In late 1978, the approach of basing policy decisions on variables other than the estimated output-gap level was consolidated when the Bank of Japan announced forecasts for M2 growth (see, e.g., Hamada 1985). These were not formally labeled targets, but Bank of Japan Governor Morinaga said, "I'll carry out monetary policy while closely watching the movement of money supply" (*NYT*, December 28, 1978). Thus, even before the series of discount-rate increases began in 1979, Morinaga had signaled that high money growth would lead the authorities to raise interest rates.

Not only were the authorities tightening ahead of much of the actual increase in CPI inflation, but they were still ahead of much opinion on the role of monetary policy in fighting inflation. A May 1979 commentary in the Kyodo News Service on the Bank of Japan's

[20]See section 4.5 below for further discussion.

discount-rate increase was entitled "Credit Policy Unlikely to Affect Inflation." It said that strengthening demand had "not as yet been much of a factor" in wholesale price increases and downplayed the contribution monetary policy could make if CPI inflation did worsen: "[R]eliance must be put on other policy measures. Fiscal policy will have to be tightened.... More vigorous action must be taken to strengthen the yen's exchange rate and bring down the cost of imports. Imports, especially of manufactured goods, must be increased.... The means to check inflation are readily available" (*KYO*, May 23, 1979).

Further discount-rate increases took place over 1979–80, with a 100-basis-point increase in February 1980 justified by Bank of Japan Governor Mayekawa as an inflation-containing action in the face of a tightening of the demand-supply position (*JT*, February 19, 1980). In contrast to the tentative support for incomes policies voiced by policymakers during 1970 and 1974, a Bank of Japan official observed the following in 1980: "No incomes policy could conceivably be effective.... [O]rthodox policies ... are valid enough to check home-made inflation."[21]

4.5 Lessons from Japan's Experience

Several lessons emerge from Japan's Great Inflation and disinflation. First, while Milton Friedman's (1990, 107) observation that "no control of individual prices nor of individual wages" occurred during the 1973–74 monetary tightening is an overstatement—Japan's officials did impose limits on price increases on specific goods—the controls applied to a small portion of the price index, and inflationary pressure suppressed by the controls could easily be transferred to other prices. Therefore, Japan's disinflation cannot plausibly be attributed to incomes policy.

Second, Japan, like other countries, had a period during which policymakers were inclined to adopt general wage and price controls because they believed inflation had become cost-push in character. But in Japan's case the strongest doubts came just before a drastic decline in inflation in 1974–75, and this decline, coming in the wake

[21]Suzuki (1981, 412).

of severe monetary restraint, served as a powerful rebuttal of the view that inflation was insensitive to negative output gaps.

Third, the recognition by Japan's policymakers of an excess-demand problem in 1973 was superior to that of outside agencies such as the OECD (1973). This may have reflected greater weight given to money (M2) growth, which had risen about 8 points over 1970–73, as an indicator of pressure on aggregate demand.

Fourth, the post-1973 slowdown meant that Japanese policymakers' output-gap estimates were probably substantially biased in the later 1970s; nevertheless, disinflation proceeded over these years. Again, the emphasis on money growth may have helped in reducing the weight given to gap estimates. But Japan's officials also indicated that they tightened monetary policy in response to positive output-gap *growth.* This "speed-limit" policy reaction could be rationalized by a variety of theories. The most traditional, but least consistent with modern inflation analysis, is that the growth rate of the gap appears directly in the Phillips curve. More consistent with a New Keynesian analysis is the view that output-gap changes suggest revisions to the expected path of the output-gap level, and so a revised inflation forecast. Responding to them is in effect a backdoor way of responding to the correctly measured (expected path of the) level of the output gap. Regardless of the specific rationale, emphasis on output-gap growth insulates monetary policy decisions from output-gap mismeasurement, since errors in level estimates tend to be persistent and largely cancel from the growth rate, as discussed in Giannoni (2002), Orphanides (2003), and Orphanides and Williams (2006).

In the present instance, there was a specific type of output-gap error that policymakers' response to growth rates protected against: the actual late-1970s output gap in Japan was closed well before real-time estimates of the gap stopped being negative. Percentage changes in variables have turning points that precede those of their corresponding levels,[22] so a policy tightening in response to an *estimated* positive output-gap *growth rate* proved, ex post, to be a tightening in response to an *actual* positive output-gap *level.*

[22]See, e.g., Culbertson (1960) for a vintage discussion.

5. Germany

This section covers the inflation-disinflation experience in Germany over 1969–80. I draw on coverage of, and commentary on, monetary policy and inflation in the German press; similar discussions of Germany in several other countries' newspapers; and Bundesbank reports, speeches, and testimony.

5.1 1969–71: Orthodox Response to Inflationary Pressure

Germany was more integrated than Japan into the international financial system over the Bretton Woods period, and so its policymakers' scope to vary the official discount rate for reasons other than the exchange rate parity was more limited than their counterparts' in Japan. Beginning in the late 1960s, however, the authorities began taking steps to shield Germany from U.S. monetary expansion. These measures included temporary exchange rate floats in late 1969 and May–December 1971, and an intensification of foreign-exchange controls in mid-1972, discussed below.

The limited monetary policy independence bought by these measures led to a series of monetary tightenings, including an increase in the discount rate from 6 percent to 7.5 percent during March 1970. Economic overheating and inflationary pressure were cited as reasons for the tightening, with the Bundesbank's vice president, Otmar Emminger, observing, "If there is no improvement in wage and price developments, we'll naturally be forced to go on making monetary policy very restrictive" (*DS*, July 20, 1970, *a.t.*). By the time this statement was published, however, the constraints imposed by the exchange rate policy had been made plain by the Bundesbank's removal of some of its March tightening, with the discount rate cut from 7.5 percent to 7 percent in July (*WSJ*, July 16, 1970). With its room to move on interest rates limited, the Bundesbank attempted to rely on reserve-requirement increases in April and June 1970 (*JT*, June 20, 1970), measures unlikely to affect aggregate demand when not accompanied by interest rate increases.

A speech by the Finance Minister, Karl Schiller, in September 1970 clearly recognized the absence of a long-run inflation/unemployment trade-off: "Inflation is like a drug. For a short time it makes our society feel 'high'.... Then it becomes apparent that the 'trip' has not solved any problems, and even created new ones." Schiller was, however, eclectic in his picture of the solution to inflation, describing incomes policy, monetary policy and fiscal policy as "complementary sets of policy instruments" (*AUP*, September 24, 1970).

5.2 1971–72: The Monetary-Policy-Neglect Period

It was from early 1971 until mid-1972, in the wake of double-digit annual nominal-wage growth in 1970–71 and the apparent elimination of excess demand, that cost-push views reached their high-water mark in German debate. As early as January 1971, an authoritative statement making heavy concessions to the cost-push view appeared in the form of an address by Bundesbank President Karl Klasen. In line with cost-push views, Klasen endorsed the position that present wage-growth rates were "not justified by economic conditions," excess demand having passed. Klasen saw the German government's "Concerted Action" program—a consultation process among labor, firm, financial, and government leaders—as a potentially valuable anti-inflation instrument. Nevertheless, Klasen demonstrated that thinking in Germany at this point was more orthodox than elsewhere. In contrast to Federal Reserve Chairman Arthur Burns's position at the time, Klasen deplored compulsory price controls as part of the solution; moreover, he rejected the view that monetary policy actions had become ineffective, and indeed indicated that the Bundesbank's recent decision not to cut the discount rate reflected a wish "to prevent continuing price rises" (*AUP*, February 2, 1971).

In the same spirit, continuing inflation (at about 4.5 percent per annum) and the danger that monetary easing could help transfer wage pressure into price inflation were cited by the Bundesbank when it again held the discount rate constant in March (*IP*, March 18, 1971). This position received support from a front-page editorial entitled "Cost Inflation" in the *Frankfurter Allgemeine Zeitung*. Proclaiming that it was "unambiguous how strongly the wage spiral has contributed to the decline in the value of money," the newspaper rejected the view of "some critics ... that the Bundesbank's measures are not effective" and supported monetary restraint as a response to wage push (*FAZ*, March 18, 1971, *a.t.*).

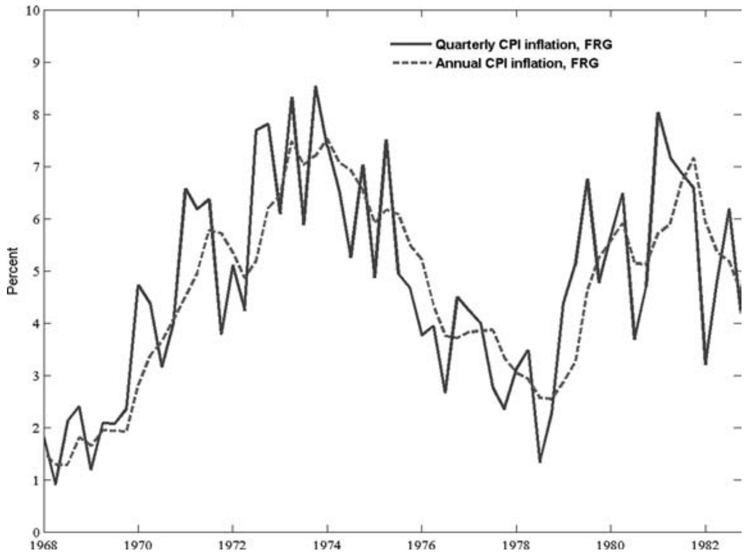
This position was in harmony with the monetary view of inflation, which states that monetary policy can block current wage developments from spilling over into expectations of future price inflation.

But it was over precisely this period, early 1971, that a shift took place in official opinion away from orthodox views on inflation behavior. The motivation for this shift was the change in the aggregate-demand picture. Estimates by the OECD constructed in early 1973 (OECD 1973, 29–32) suggested that Germany's output gap peaked at +1.3 percent in the first half of 1970, averaged +1.0 percent in the first half of 1971, and turned to −1.5 percent in the second half of 1971. These were not based on German government sources, but the January 1971 statement by Bundesbank President Klasen mentioned above indicates that German policymakers thought the gap was 0 by 1971:Q1. Therefore, like the OECD, German policymakers thought that excess demand had peaked in 1970, but unlike the OECD, they believed it had dissipated by early 1971.

With inflation rising during 1970 and 1971, policymakers and opinion leaders were susceptible to explanations for inflation that attributed it to nonmonetary factors—instead of (correctly) viewing the inflation as the result of the pre-1970 monetary ease working its way through the system. As we have seen, by early 1971 Klasen was already attributing substantial wage inflation to cost-push forces. Therefore, the longer price inflation proceeded at high rates after 1970, the more likely it was to be attributed solely to factors other than excess demand. Sure enough, Jürgen Husmann, the deputy economics director of the German Employers' Federation, said in March 1971, "[T]he recent price increases are not demand-induced. They arise from increases in wage costs" (*DT*, April 1, 1971). The *Financial Times* had already been quick off the mark, stating in January 1971 that "the easing of demand pressure has had no effect on the inflationary wages-prices spiral. . . . German inflation is now of the cost-push rather than the demand-pull variety, and consequently more difficult to control" (*FT*, January 1, 1971).

Assessments like this prevailed in policy circles too, with Finance Minister Schiller in September 1971 describing "cost pressures" as a separate cause of inflation distinct from excess demand (*AUP*, October 5, 1971). Indeed, shortly after the May 1971 exchange rate

Figure 3. CPI Inflation in the Federal Republic of Germany (FRG)



float, Schiller saw the economy as having featured the removal of overemployment, and he believed "a new chapter" had begun, with incomes policy (the Concerted Action) playing a large role in controlling inflation (*DW*, June 23, 1971, *a.t.*).

This "new chapter" indeed featured distinct policies from 1970, including looser monetary policy. The dissipation of monetary restraint was evident in a 2-percentage-point decrease in the discount rate over calendar-year 1971; this occurred despite annual CPI inflation *rising* about 2 points during the year (see figure 3). The interest rate cuts continued over the May–December 1971 interval, which the Bundesbank considered a floating-mark period (*BBAR*, April 1972, 18), so it is difficult to argue that these cuts reflected policymakers balancing inflation control against exchange rate stability. Rather, the belief that inflation could be solved by incomes policy implied that monetary policy was free to stimulate the seemingly weak economy. Overambitious goals for incomes policy had been made public in January 1971 by Bundesbank President Klasen, who had voiced the opinion that Concerted Action offered the prospect of moving from a situation where "[union] members judge the success of the trade union leadership on the success of the wage struggle" to "a higher level" where "the performance should be judged on trade unions' part in the general economic report of the Federal Government" (*DW*, January 23, 1971, *a.t.*).

Finance Minister Schiller resigned in mid-1972 in protest against new policy decisions, including the adoption of exchange controls (Fels 1977, 613). Since Schiller was arguing for a freer financial system, it would be tempting to conclude that his defeat was a further setback for orthodox macroeconomic policies. That would not, however, be an altogether appropriate conclusion, since Schiller's departure increased the Bundesbank's ability to carry out monetary tightening. Schiller favored a fixed exchange rate with minimal exchange rate controls; but if monetary restraint was to commence under fixed rates, Germany needed foreign-exchange controls. With these controls imposed, the German authorities were indeed able to raise interest rates in late 1972. From a monetary-control perspective, the reimposition of exchange controls was desirable. The policy package involving foreign-exchange controls was thus the lesser of two evils, even though less appropriate than the alternative, not-yet-on-the-table option of a permanent float, monetary restraint, and no exchange controls.

Schiller's departure was also beneficial in promoting a shift away from incomes policy as an anti-inflation weapon. In August 1972, shortly after assuming office as Finance Minister, Helmut Schmidt was asked to comment on the United States' wage and price controls. His response was to cast doubt on their value: "I think a short-term success of the U.S. experiment is quite possible. But in the long run controls create serious distortions of market forces. You've got to grasp how wedded we are to the concept of free markets and free competition as a foundation of a productive economy. We consider the basically free play of prices on the market and the absolute independence of employer and employee in wage negotiations as vital to this process. I don't really think much of trying to interfere in this autonomous play of forces with such things as wage and price controls" (*NW*, August 21, 1972). Schmidt's views stand in contrast to the enthusiastic, even utopian, prior statements by Schiller and Klasen to the effect that incomes policy opened up a new chapter in which unions' objective functions were transformed.

The late-1972 interest rate increases were preemptive in the sense that they amounted to tightening while the output gap was still believed to be negative—the Bundesbank's report on 1972 gave it as −1.5 percent on average (*BBAR*, April 1973, 11). The 1972 tightening was in response to rapid money growth and to the signal this gave of prospective overheating in 1973. Thus monetary policy in 1972 focused on monetary aggregates but did so with a preemptive outlook regarding aggregate-demand developments. There is therefore considerable continuity from late-1972 policy to the late-1974 policymaking described by Bernanke et al. (1999, 47), which they cite as notable for preemption (though a preemptive easing, as opposed to the preemptive tightening of 1972).

5.3 1973–75: New Regime

The transfer of inflation control to monetary policy, already underpinning the 1972 discount-rate increases, was boosted further in early 1973, notably in a call by the federal government, documented by von Hagen (1999a, 683), for monetary policy to carry out disinflation. While von Hagen emphasizes that this preference for monetary policy reflected the executive branch's reluctance to tighten fiscal policy, for the present discussion, two other very important aspects of the government's call are crucial and put it in a more favorable light. First, it was an affirmation at a high policymaking level that restricting aggregate demand was essential for reducing inflation. This contrasts with, e.g., the United Kingdom and Ireland at the time, with their governments' belief that monetary stimulus actually reduced inflationary pressure. Second, a disinflationary monetary policy was acknowledged as compatible with lack of fiscal consolidation. This contrasts with the position of Arthur Burns in the United States that fiscal deficits automatically pushed up money growth and/or inflation.

During the second quarter of 1973, the Bundesbank raised the discount rate in two steps to 7 percent. At a press conference following the first of these increases, Bundesbank President Klasen affirmed that the increase was designed to fight inflation. He did, however, indicate he might support a wage-price freeze if inflation deteriorated, adding that "we must use all other available measures" before contemplating that step (*JT*, May 5, 1973).

A firmer rejection of controls as a solution to Germany's inflation took place the following week, when the Brandt government announced its own anti-inflationary measures, concentrating on fiscal restriction and excluding wage and price controls (*SCMP*, May 11, 1973).[23] Commenting on the measures, the "father" of Germany's postwar currency, Ludwig Erhard, applauded the government's rejection of wage-price controls, adding, "I think perhaps we now should have another currency reform in miniature" (*JT*, June 22, 1973).

In fact, as he spoke, Germany was undergoing a "currency reform in miniature"—in the form of the regime of a floating exchange rate and inflation-oriented monetary policy, started in 1973. Germany's regime change is frequently dated to the announcement of monetary targets in December 1974, a dating also implicit in the estimation of reaction functions on samples beginning in 1975 at the earliest.[24] But money growth had been cited as a consideration when the Bundesbank began tightening in October 1972 (*JT*, October 8, 1972). By October 1973, Helmut Schlesinger, a Bundesbank director, was describing March 1973 as the date when "[m]onetary policy ... moved back into the center of anti-cyclical policy" (*FT*, October 8, 1973). And, indeed, the Bundesbank in 1975 described its 1974 strategy as "to continue the fight begun in 1973" (*BBAR*, April 1975, 14), and though monetary targeting was not introduced until the end of 1974, a Bundesbank official testified in 1980 that the Bundesbank "for internal purposes had already established a target for its own orientation for the year 1974" (Dudler 1980, 299).

[23] Chancellor Brandt reaffirmed the rejection of controls in a television interview in February 1974: "Such steps have not proved to be effective. In our neighboring countries, such steps did not bring about any better results" (quoted in *JT*, February 3, 1974).

[24] See, e.g., Bernanke and Mihov (1997). As discussed below, Bernanke et al. (1999, 43–47) do offer official statements that they suggest are evidence of the influence of monetarism on German policy thinking in the period predating official monetary targeting. But the examples they cite, as well as the quotation offered by von Hagen (1999a, 690), are not in fact unambiguous evidence on this point. They could imply that monetary restraint is merely necessary for (or "compatible with," to use the words that von Hagen quotes) inflation control, whereas monetarism treats monetary restraint as necessary *and sufficient* for inflation control. The material I present is, by contrast, more clear-cut in showing that key Bundesbank personnel embraced the monetary view of inflation.

Therefore, the regime in force from December 1974 was a formalization of that prevailing throughout 1974, itself in turn a continuation of the regime begun in March 1973. Indeed, most of the "heavy lifting" carried out during the regime—substantial interest rate increases—had been done by the time monetary targeting was formally introduced, with the discount rate being cut from October 1974, annual CPI inflation having peaked (at 7.5 percent) in 1974:Q1.

What were the doctrinal changes underlying the regime change? The price-stability task of monetary policy was already enshrined in the Deutsche Bundesbank Act of 1957.[25] But to regard price stability as the primary task of monetary policy is only part of the way to a modern view of inflation control. What is also needed is recognition that nonmonetary instruments such as incomes policy are redundant and ineffective as means of fighting inflation. Here, changes in personnel at the top of the federal government were an important element, with Finance Minister (from May 1974, Chancellor) Schmidt more inclined than his predecessor to focus on monetary policy and to downplay incomes policy.[26]

As far as the Bundesbank is concerned, however, there is evidence that President Klasen, while supportive of tightening money in 1973, did *not* undergo a profound change in his views relative to 1971. At that time, he had subscribed to a mixed cost-push/monetary view of inflation, with his behavior from mid-1971 suggesting a move to a harder-line cost-push view, giving up on the idea that monetary actions could rein in wage inflation; further, as noted above, in mid-1973 he had withdrawn his 1971 opposition to compulsory wage and price controls (though the government did not adopt them after all). The monetary policy tightening in 1972–73 was compatible with this position, as it was a reaction to a prospective positive output gap and did not in itself signify a denial of the importance of cost-push factors. Even in June 1974 President Klasen described the monetary tightening as amounting to "[w]hat could be done by monetary means" against inflation, with low inflation also requiring improved attitudes about sharing the economic "cake" (*AB*,

[25]See, e.g., Deutsche Bundesbank (1980, 291).

[26]The Bundesbank's Otmar Emminger subsequently noted that Schmidt's arrival heralded a "shake-up . . . [of] great significance" (Emminger 1977, 34).

June 27, 1974). Klasen's nonconversion to the monetary view of inflation was also reflected in the Bundesbank's most authoritative statements, such as the 1972 *Annual Report*, which stated the following on its opening page: "Monetary policy alone cannot avert the danger of inflationary expectations gaining strength" (*BBAR*, April 1973, 1).[27]

A sharper change in attitude to inflation control occurred among other directors of the Bundesbank. Chief Economist Helmut Schlesinger was much more influenced by monetarism, citing "theoretical arguments and empirical findings produced largely in the U.S." Schlesinger noted that a monetary approach to inflation control was "nothing new from the German point of view" (*FT*, October 8, 1973), but the new research on the subject helped convince Schlesinger that German policymakers had been wrong to look to devices other than monetary policy for inflation control.

Bundesbank Vice President Emminger, like Schlesinger, took a stronger view than President Klasen on the contribution that monetary policy could make. Emminger affirmed that the Bundesbank could deliver price stability, and he spoke out against critics. "For example," he observed in November 1973, "people say: 'Anti-inflation policy, well and good; but in the cost of living there's a large proportion of administered prices which are insulated from market laws and thus from the overall instruments of fiscal and monetary policy.' According to this argument, no matter how hard we try, we can never get below a given bedrock inflation rate determined

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[27]This quotation puts in perspective the statement in the same report that a "persistent and accelerating decline in the value of money is impossible without a corresponding [monetary] expansion . . . [and] the monetary sphere in its own right not infrequently promotes the inflation of prices and wages" (*BBAR*, April 1973, 24). Bernanke et al. (1999, 50) interpret this quotation as implying that "monetarism was having a significant impact on policymaking inside the Bundesbank." This conclusion is accurate—see the October 1973 quote from Helmut Schlesinger given here—but does not follow from the quotation, which is a weaker statement than what is implied by a monetary view of inflation. To say that "accelerating" inflation requires monetary accommodation, and that monetary expansion "not infrequently" promotes inflation, could simply be acknowledging that monetary expansion can create excessive demand and compound cost-push inflation. Unlike later statements by the Bundesbank, quoted below, this statement does not reject the possibility that cost-push forces can be a source of maintained inflation, nor does it acknowledge that monetary restraint is a sufficient condition for price stability.

by the administered prices, let us say 5%." Emminger rejected this "defeatism about the value of monetary policy." The proportion of prices and wages that responded to market forces was large, and he noted that even administered prices ultimately responded to market forces (*AUP*, November 9, 1973, *a.t.*).

It is therefore appropriate to conclude that the monetary regime that began in 1973 was influenced by the monetary view of inflation, but President Klasen underwent a less significant change in opinion than his subordinates. To emphasize, this did not amount to a difference among policymakers on the costs of inflation; Emminger and Schlesinger shared the goal of price stability with Klasen. Rather, it amounted to a different perception of what had created price stability before 1970 and what had changed since 1970. Klasen was more concerned with reinforcing specific institutional arrangements that had existed before 1970 such as Concerted Action, believing the need for incomes policy to have become more important in the 1970s. Emminger and Schlesinger had far less attachment to the pre-1970 institutional arrangements, be they Concerted Action or fixed exchange rates. Their reasoning was that these features would not secure the truly important element that had been lost since 1970—namely, monetary stability. They saw no merit in fixed exchange rates per se, notwithstanding fixed rates being a feature of the earlier price-stability period; monetary stability could be delivered by domestic monetary restraint combined with floating exchange rates, and fixed rates could be an impediment to restoration of monetary stability. Klasen thought monetary restraint was one condition for price stability; Schlesinger and Emminger thought it was a necessary and sufficient condition.[28]

In 1973, however, with unanimity that excess demand was the immediate problem, doctrinal disagreements were less likely to manifest themselves as policy disagreements. This unanimity, and the

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[28]This conjecture about Emminger's and Schlesinger's 1973 positions is compatible with their later reputation as dissenters in post-1974 Bundesbank deliberations, as indicated by von Hagen's (1999a, 690) discussion of the Bundesbank General Council's minutes. Indirect support for associating Klasen with the nonmonetary approach to inflation is also provided by the fact that neither von Hagen's (1999a) nor Bernanke et al.'s (1999) account of Bundesbank monetary targeting mentions Klasen, despite monetary targeting being inaugurated during his presidency.

solidarity between the government and the Bundesbank on the 1973 policy change, proved important, as the new course was criticized by many leading financial commentators. Jürgen Ponto, president of Dresdner Bank (the Federal Republic's second-largest commercial bank), said in November 1973, "The doubts whether a lasting stabilization of cost and prices can be achieved by holding fast to the current restrictive credit and fiscal policy course are being only reinforced." Ponto claimed that interest rate increases were having a cost-push effect "rather than the desired anti-inflationary effect" (*JC*, November 5, 1973). *The Economist*, too, was extraordinarily critical. In June 1973 it asserted: "Chances for an early dip in cost-push inflation therefore look slim.... [T]ight money ... will merely turn boomflation into stagflation" (*TE*, June 9, 1973). In November it editorialized that "the right economic recipe is re-expansion and an incomes policy" and judged that the government's rejection of controls and its embrace of demand restraint "could hardly be a worse policy for Germany at this time" (*TE*, November 10, 1973).

5.4 1975–80: Consolidating the New Regime

In the second half of the 1970s, the regime change was consolidated by data outcomes—low inflation—and personnel changes—notably the accession (in 1977) of Otmar Emminger to the Bundesbank presidency. An interview Emminger gave in early 1975, while still Bundesbank vice president, was notable for his emphasis on aggregate demand control. He stressed that Germany had gotten "inflation under control ... [by] apply[ing] the classical medicine of restrictive fiscal and monetary policy." Asked if Germany's monetary policy solution could be applied to the United Kingdom, Emminger expressed the reservation that "[w]hen you already have very high and firmly established inflationary expectations, it is difficult to break them with restrictive fiscal and monetary policy alone" (*TG*, March 5, 1975). This phrasing of the issue was probably a diplomatic way of avoiding direct criticism of UK policymakers; Emminger was being interviewed by a British newspaper. It is uncontroversial, according to the monetary view of inflation, that monetary policy might not reduce inflationary *expectations* at the same speed that it removes inflationary pressure.[29] This does not prevent monetary policy from reducing inflation, but it increases the short-term output costs of a disinflation. Incomes policy is sometimes invoked as helpful in these circumstances, by providing a direct link between nominal contract arrangements and the government's disinflation program. The problem with using this as a defense of incomes policies pursued by countries like the United Kingdom in the 1970s is that incomes policy was seen in those countries as an anti-inflation policy in itself, a view Emminger rejected: "it is nearly impossible to break established inflation by relying on incomes policy alone" (*TG*, March 5, 1975).

Later in 1975, Emminger was more outspoken about nonmonetary approaches to inflation control. Emminger referred to the "baffling complexities" of incomes policies, and went on as follows: "[W]hatever the initial causes of a particular price inflation—they may be entirely exogenous like bad harvests, the oil price increase, etc.—in the longer run price inflation can continue only if it is accommodated by permissive monetary policies. Inflation is a monetary phenomenon. Thus the responsibility of the central banker is always involved" (*AUP*, December 19, 1975).

Similarly, in 1977 the Bundesbank's Helmut Schlesinger said, "In the medium run, general price increases cannot occur without excessive expansion of the money stock" (*AUP*, October 11, 1977). This firmness contrasts with positions of other countries' policymakers at the time, not only elsewhere in Europe but also in the United States. Federal Reserve Chairman Arthur Burns, in one notable statement, said he was confident that monetary policy could prevent inflation only "if private enterprise doesn't go wild and if Congress stops legislating inflation"[30]—so that in contrast to the Bundesbank orthodoxy, monetary restraint in Burns's view was merely one of many conditions for inflation control.

[29] That monetary policy can ultimately—whatever short-run inertia exists in inflationary expectations—pin down those expectations fully by keeping the expected output-gap path at zero, distinguishes the monetary view of inflation.

[30] Arthur Burns, November 9, 1977 testimony, in Committee on Banking, Housing, and Urban Affairs (1977, 26).

Several practical features of the new monetary regime emerged.[31] First, the Bundesbank did not accommodate the 1973–74 oil shock. President Klasen said in 1974 that it "would be wrong to inhibit indispensable adjustment processes by artificially enlarging aggregate demand" (*AB*, June 27, 1974), while in the December 1975 quotation above Emminger stated the non-accommodation principle. Quarterly inflation peaked at end-1973 and annual inflation in 1974:Q1, well ahead of the peaks in other countries. For Germany, the peak reflected a one-time price-level jump from the OPEC shock, as well as pre-1973 monetary ease, and did not reflect accommodation of the oil shock.[32]

Second, while focusing on money growth, the authorities did not discard evidence from real variables; as noted above, in 1973, Bundesbank Director Schlesinger had described the new policy as "anticyclical," while the Bundesbank's annual reports in the 1970s and into the 1980s plotted estimates of potential output (e.g., *BBAR*, April 1981, 11).[33] The attention to the output gap in policymaking did not contradict the reaffirmed orientation of monetary policy on inflation control. But it did raise the possibility that output-gap mismeasurement, occurring especially with the post-1973 economic slowdown, would provoke inappropriate monetary easings, as Orphanides (2003) argues occurred in the 1970s in the United States. The Bundesbank partially avoided this problem by promptly recognizing some of the post-1973 slowdown. Emminger said in 1975 that policymakers "definitely" expected permanently lower economic growth because of slower growth in the labor force and other structural changes (*TG*, March 5, 1975), as well as "a somewhat higher level of unemployment than we have been used to since the 1960s" (*CAP*, February 1975, *a.t.*).

[31]Tactical features of the regime, i.e., the operating procedures used by the Bundesbank in the financial markets, are not discussed here due to my focus on strategy. Bernanke et al. (1999), Issing (1997), and von Hagen (1999a) provide extensive discussions of Bundesbank operating procedures.

[32]Therefore, the suggestion by Clarida and Gertler (1997, 375) that the Bundesbank heavily accommodated the first oil shock does not seem to be warranted.

[33]Von Hagen (1999b, 434) and Gerberding, Seitz, and Worms (2005) also indicate that countercyclical considerations weighed heavily in the Bundesbank's internal deliberations.

Third, German policymakers, like their Japanese counterparts, had "speed-limit" concerns about inflationary pressure. For example, in December 1975 Emminger observed that "an economy may run into bottlenecks long before reaching full employment" (*AUP*, December 19, 1975, 3), a position he reaffirmed in 1978 (*AUP*, February 3, 1978). Consistent with this concern, Bundesbank Director Schlesinger gave "current utilization of the production potential, and possible changes in this utilization" as factors that affected the choice of each year's money-growth target (*AUP*, October 11, 1977, 3). Of these, the change in utilization evidently came to have the more systematic effect on policy decisions, at least after 1978: for the twenty years beginning in 1979:Q1, Gerberding, Seitz, and Worms (2005) find that responses to the output gap in the Bundesbank's policy rule are small and statistically insignificant, while responses to output-gap growth are significant. Mismeasurement of the level of the output gap was substantial in Germany in the second half of the 1970s, as they show, so the basing of policy on output-gap growth was beneficial. The official statements cited above on the link between bottlenecks and inflation suggest that the Bundesbank's focus on output-gap growth did not arise from skepticism about the level estimates but from belief in a speed-limit term (in addition to a gap-level term) in the Phillips curve.[34]

Fourth, incomes policy did not play a part in the disinflation. As discussed in the introduction, advocates of incomes policy outside Germany attributed German inflation success to union-government cooperation regarding nominal-wage growth. Such accounts have no merit. The consultation body, Concerted Action, which had been cited as an anti-inflationary tool during the heyday of cost-push views in Germany, was disbanded in 1977.[35] A Bundesbank official explained in 1980 that the authorities "gave up" on Concerted Action, adding, "I do not think that we or the trade unions felt that this was a very important arrangement as far as actual

[34] Equally, by seeming to explain the short-run coexistence of negative output gaps and inflation, the speed-limit perspective on Phillips-curve dynamics probably slowed down German policymakers' revision of their output-gap estimates to more accurate values.

[35] Braun (1986, 240).

policymaking is concerned" (Dudler 1980, 305). On the more general question of incomes policies, the Bundesbank stated: "The Bundesbank and Federal Government have never regarded administrative wage and price controls as an alternative (or supplement) to monetary policy.... The Government neither intervenes directly in specific wage and price decisions nor attempts to hold trade unions or employers' associations to formal wage and price guidelines.... [T]here is practically no convincing evidence of the lasting success of any variant of direct income[s] policy" (Deutsche Bundesbank 1980, 295).[36]

Fifth, the Bundesbank did concede an influence of cost-push pressures on inflation in their published estimates of the amount of "unavoidable" inflation in the year ahead, i.e., the Bundesbank's "price assumption" or inflation target. This concept encompassed not only the inflationary pressure built in by prior monetary policy decisions but also price-level shocks that were conceded as having an impact effect on inflation. A Bundesbank official in 1980 gave "higher raw material prices or oil prices" as influences on the unavoidable inflation rate (Dudler 1980, 299). The impact of such factors on inflation is, however, compatible with the monetary view of inflation, since according to that view, it is expected future inflation, not current inflation, that is pinned down by monetary policy alone. In addition, the announced price assumption tended to decline, settling at a rate of 2 percent after 1985,[37] showing that even the "unavoidable" component of inflation was regarded as an endogenous variable at horizons beyond the short run.

It is likely, however, that the Bundesbank overestimated the importance of cost-push factors in the determination of unavoidable inflation. For example, in 1978 the Bundesbank believed, in the words of one official, that the mark was "faced with the prospect of an uncontrolled appreciation" (Dudler 1980, 306), and undertook

[36] Similarly, a description by Germany's federal government of economic policy said, "[W]age freezes or the fixing or limiting of wage increases are not included amongst the instruments employed in evolving the State's economic policy" (quoted in *UKPD*, November 9, 1978, p. 1218).

[37] See Coenen, Levin, and Christoffel (forthcoming) and Gerberding, Seitz, and Worms (2005).

unsterilized intervention to offset some of the upward pressure.[38] This was despite the fact that over this period, President Emminger gave "one to three percent" as the only inflation rate that the Bundesbank considered tolerable (*BKR*, September 1978). It is clear that the Bundesbank thought it could get away with monetary stimulus in 1978 despite its monetary view of inflation because it believed that the negative impact effect on inflation from mark appreciation would offset the upward pressure coming from the monetary easing. In 1980 a Bundesbank official was very open about this, revealing that the authorities had "felt that pursuing a low interest-rate policy and allowing monetary growth to accelerate would not have a detrimental effect" on inflation (Dudler 1980, 306–7). This proved not to be the case, as annual CPI inflation exceeded 5 percent in 1980, a rate above what the Bundesbank regarded as acceptable even in the face of the second oil shock.[39] An alternative policy in 1978, which disregarded the nonmonetary influences on inflation, would have led to this error being avoided. The misjudgments underlying the 1978 episode probably played a part in leading Karl-Otto Pöhl, who became Bundesbank president in 1980, to state, "Interest rates should be set according to domestic monetary conditions and the exchange rate should be left to go where it will."[40]

6. Did Policymakers Try to Exploit a Phillips-Curve Trade-Off?

The preceding sections drew attention to some lessons regarding 1970s policymaking in Japan and Germany. In particular, both countries switched from a problematic nonmonetary approach to inflation control in 1971–72 to a monetary approach to inflation control in 1973; confidence in monetary policy was reinforced by falling inflation in 1975; and the particular variables policymakers used to measure excess demand—namely, money growth and the change in the output gap—enhanced preemptiveness of policy and reduced

[38] See von Hagen (1999a, 693) for other details of this episode based on different sources.

[39] "We certainly would feel that a rate of 5% is too high . . . for what we might accept as an unavoidable structural built-in inflation element" (Dudler 1980, 301).

[40] Quoted in Thatcher (1995, 479).

the influence of estimated output-gap levels on policymaking. I now consider a further lesson about policymaking and inflation behavior that emerges from joint study of the two countries.

An important element of many accounts of the United States' Great Inflation, including those of DeLong (1997) and Sargent (1999), is the position that policymakers were guided by the view that there was a long-run, exploitable Phillips-curve trade-off between inflation and unemployment. The evidence in this paper casts serious doubt on this story as an explanation for the Great Inflation that is valid across countries. Belief in Phillips-curve trade-offs simply was not an important factor behind policy mistakes in Germany and Japan.

A belief in a long-run Phillips-curve trade-off was not the source of Germany's 1970s inflation problems. The essentials of the long-run vertical-Phillips-curve view had been voiced officially in Germany in 1970, by Finance Minister Schiller, who shared his predecessors' goal of price stability.[41] Moreover, in 1975 Bundesbank President Klasen said it was "wrong" to believe in "a long-lasting solution to unemployment through more inflation" (*NW*, February 17, 1975). Where policymakers—notably Schiller and Klasen—lapsed in the 1970s, it was in succumbing to cost-push views, not trade-off pursuits.

Their lapse also indicates that denying a long-run Phillips-curve trade-off is not enough. A sound official doctrine also needs to be subtle by taking care not to reject all aspects of Phillips-curve analysis. From the perspective of modern macroeconomics, the phenomenon of stagflation reflects the impact on inflation dynamics of two terms that appear in a correctly specified Phillips curve: expected inflation and shocks to potential output. But 1970s policymakers could be—and indeed were, especially outside Germany—tempted to interpret stagflation as instead revealing that *no* relationship existed between unemployment (and so the output gap) and inflation, or that the relationship was positive under all circumstances. Such misinterpretations, while successful in leading policymakers away from attempts to exploit trade-offs, are unhelpful because they obscure the fact that the way to remove inflation is to work

[41] As former Chancellor Erhard put it, "I am convinced the maximum rate of price increases should be 2 percent—but 1 percent is preferable. Herr Schiller wants that, too" (quoted in *KCS*, July 1, 1970).

through the aggregate-demand channel. It is not, therefore, a badge of honor to be so hawkish about inflation as to believe that inflation has only a positive relationship with unemployment. This variant of hawkishness obscures the mechanisms connecting monetary policy actions to inflation, so it is not really a road to a low-inflation regime.

In this light, a notable contribution by Helmut Schmidt to German economic policymaking, in addition to transferring inflation control to monetary policy, was to restore a balanced view of the unemployment/inflation relationship. Schmidt voiced a subtle interpretation, stating that the message of the data was “that the correlation between unemployment and inflation is different, but that there is a fundamental connection” (*DZ*, November 8, 1974, *a.t.*), and that it was “too simple to say that inflation causes unemployment” (*SZ*, June 24, 1975, *a.t.*). Similarly, Helmut Schlesinger of the Bundesbank noted that raising economic growth and cutting inflation were not compatible in the short run, because “to reduce inflation means dampening the economy” (*CT*, October 22, 1981). These calls for subtlety are in harmony with the attitude of Milton Friedman, who wrote in 1979: “Orthodox wisdom has it that unemployment is a cure for inflation. A minority has it that unemployment causes inflation. Both views are half-truths” (*NW*, November 12, 1979).

In Japan, there was no point where belief in a long-run trade-off was the official view. In February 1970, Finance Minister Fukuda gave Japan’s unavoidable inflation rate as 4–5 percent (*JT*, February 20, 1970), with the government stating that “to maintain our economic growth, some degree of price increase is inevitable” (*JT*, March 3, 1970). This claim was not, however, based on a Phillips-curve trade-off calculation; the Phillips-curve trade-off implies that higher inflation can buy a higher *level* of output, whereas the government’s statement referred to the inevitability that moderate inflation would coexist with steady-state economic *growth*.

It is likely that the Japanese government was not simply stating that reducing inflation below 4 percent would require a temporary disruption of growth. Rather, its references to 4–5 percent inflation as unavoidable or inevitable probably indicated a view that superneutrality violations (e.g., “wheel-greasing” effects of inflation) existed that made 4–5 percent inflation rates (approximately the rates observed during the 1960s) conducive to continuation of Japan's 1960s economic growth.[42] Certainly higher inflation rates were ruled out: even in 1970, the government regarded bringing inflation back below 5 percent as a desirable immediate goal and reducing it to 3–4 percent as a long-term goal (*JT*, February 20, 1970; March 3, 1970). The subsequent 20-point rise in Japan's inflation rate cannot plausibly be attributed to government exploitation of a trade-off calculation: as noted, the government wanted to bring inflation back below 5 percent, while in 1971–72 it had nonmonetary (and therefore non-Phillips-curve) views of the inflation process. After 1973, Japanese policymakers indicated that they viewed inflation dynamics in terms of a conventional, long-run vertical Phillips curve.[43]

7. Conclusion

Many theories about why countries inflate take for granted that policymakers understand that inflation is a monetary phenomenon. The evidence presented in this paper suggests that these theories do not have merit in understanding Germany's and Japan's 1970s experiences. Two particular factors often cited as important in accounting for inflation outcomes—(i) government pressure on central banks to inflate and (ii) policymakers' belief in a long-run unemployment/inflation trade-off—do not appear important in understanding these countries' inflation-disinflation pattern. The suggestion that central bank independence is an important factor in delivering low inflation is belied by these countries' experiences. Japan's central bank was not independent, yet Japan disinflated early; and in Germany's case, pressure for disinflation came from Finance Minister Helmut Schmidt who, despite the Bundesbank's official

[42]For example, an empirical regularity of the 1960s was that Japan's wholesale price index was stable even as the CPI rose (Komiya and Suzuki 1977, 306). It may have been thought that the resulting relative-price pattern was one condition for Japan's steady-state growth and might be disturbed if CPI inflation proceeded at zero or very low rates.

[43]The particular favored variant of the Phillips curve in policy circles was one with stickiness in prices and flexibility in nominal wages (see, e.g., Suzuki 1985). Ball (1994, 174) and Taylor (1989, 137–42) likewise suggest that nominal-wage flexibility may be a reasonable approximation for Japan.

independence, played a major role in retrieving order in monetary arrangements after the shambles of the 1971–72 period. The argument that policymakers' belief in a long-run Phillips-curve trade-off, and subsequent acceptance of no long-run trade-off, drove monetary policy developments is also not supported for either country.

What appears necessary for a successful explanation for the Great Inflation across countries is an account that does not take for granted policymakers' understanding of the monetary character of inflation. The monetary-policy-neglect hypothesis suggests that high-inflation episodes in the 1970s reflect neither conscious acceptance of inflationary policies by governments nor denial by policymakers of the costs of inflation. The analysis of Germany and Japan in this paper suggests that this hypothesis is useful for understanding their early embrace of disinflation. Germany's and Japan's experiences in the 1970s indicate that once inflation is accepted by policymakers as a monetary phenomenon, the main obstacle to price stability has been overcome.

Appendix 1. Abbreviations for Periodicals Cited in Text

AB—American Banker (United States)
AUP—Auszüge aus Presseartikeln (Deutsche Bundesbank, Frankfurt)
AZR—Arizona Republic (Phoenix, AZ)
BBAR—Report of the Deutsche Bundesbank (Annual Report) (Frankfurt)
BKR—The Banker (London)
CAP—Capital (Germany)
CT—Chicago Tribune (Chicago, IL)
DS—Der Spiegel (Hamburg)
DT—Daily Telegraph (London)
DW—Die Welt (Hamburg)
DZ—Die Zeit (Hamburg)
FAZ—Frankfurter Allgemeine Zeitung (Frankfurt am Main)
FT—Financial Times (London)
IP—Irish Press (Dublin)
JC—Journal of Commerce (Newark, NJ)
JT—Japan Times (Tokyo)
KCS—Kansas City Star (Kansas City, MO)

KYO—Kyodo News Service (Tokyo)
NKS—*Nihon Keizai Shimbun* (Tokyo)
NW—*Newsweek* (United States and Europe)
NYT—*New York Times* (New York)
SCMP—*South China Morning Post* (Hong Kong)
STR—*Straits Times* (Singapore)
SUN—*The Sun* (London)
SZ—*Süddeutsche Zeitung* (Munich)
TE—*The Economist* (London)
TG—*The Guardian* (London and Manchester)
UKPD—*U.K. Parliamentary Debates: House of Commons* (London)
WSJ—*Wall Street Journal* (New York)

Appendix 2. Chronological List of Periodical Articles Cited on Great Inflation in Japan and Germany

Japan

Auszüge aus Presseartikeln, "Ansprache des japanischen Finanzministers Takeo Fukuda," (excerpts from speech at World Bank, September 30, 1969) October 7, 1969, pp. 12–13.

Japan Times, "Text of Sato's Policy Speech," December 2, 1969, p. 12.

Editorial. "Nikkeiren's Wage Proposal." *Japan Times*, February 1, 1970, p. 12.

Japan Times, "Fukuda Says 4–5% Rise in Prices Inevitable," February 20, 1970, pp. 1 and 5.

Editorial. "Your Yen's Worth." *Japan Times*, March 3, 1970, p. 12.

Porter, S. "Japanese Worker Saving-est in World." *Arizona Republic*, May 13, 1970, p. F8.

Kyodo News Service. "Incomes Policy Gets Debated at Seminar." *Japan Times*, July 10, 1970, p. 10.

Japan Times, "Bank of Japan Reduces Official Discount Rate by 0.25% to 6%: Policy Body Thinks Tight Money Has Achieved Purpose," October 28, 1970, pp. 1 and 5.

Japan Times, "Consumer Prices Rising: EPA Warns Against Large Wage Hikes," December 5, 1970, p. 13.

Japan Times, "Government Freezes Most Public Charges for Indefinite Period," December 10, 1970, pp. 1 and 5.

South China Morning Post, "Japan May Adopt 'Incomes' Policy," December 12, 1970, Business News, p. I.

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Editorial. "The Fiscal 1972 Budget." *Japan Times*, January 14, 1972, p. 14.

Hara, S. "World Economy Sluggish But Prospects Improving; Nixon's Plan, Surtax, Revaluation Hit Japan Hard." *Japan Times*, May 31, 1972, p. 11.

Japan Times, "Credit Squeeze Likely in Bid to Curb Inflation," December 25, 1972, p. 12.

Japan Times, "Economy Overheated, Bank of Japan Says," February 8, 1973, p. 11.

Nihon Keizai Shimbun, "Tight Money Policy," April 11, 1973 (translation in *Japan Times*, April 12, 1973, p. 14).

Japan Times, "Govt. Plans to Expand Import Quotas by 30% Over Levels for 1972," April 14, 1973, pp. 1 and 5.

Japan Times, "Govt. Ministers Agree on Need to Stabilize Prices," May 26, 1973, p. 1.

Japan Times, "Tight Money Steps Viewed Skeptically," *Japan Times*, May 30, 1973, p. 10.

Editorial. "Discount Rate Raise." *Nihon Keizai Shimbun*, May 30, 1973 (translation in *Japan Times*, May 31, 1973, p. 14).

Japan Times, "Premier Feels Govt. Tight Money Policy Will Curb Inflation," July 6, 1973, pp. 1 and 5.

Japan Times, "Tight Money Policy to Continue: Premier," November 2, 1973, p. 12.

Japan Times, "Premier Rules Out Need for Incomes Policy," December 9, 1973, pp. 1 and 3.

Japan Times, "Govt. Forecasts GNP Rise of 2.5% for 1974, Lowest in 20 Years," December 22, 1973*a*, pp. 1 and 3.

Japan Times, "Bank Rate Raised by 2% to Reach Record 9%," December 22, 1973*b*, p. 1.

Editorial. "Changed Economic Fortunes." *Japan Times*, December 30, 1973, p. 8.

The Economist, "Japanese Budget: Austerity, Gift-Wrapped," January 5, 1974, p. 72.

Japan Times, "Tanaka Requests Labor to Show Restraint in New Wage Raise Bid," February 22, 1974, pp. 1 and 4.

Japan Times, "Japan Will Slow Down Growth Rate: Fukuda," May 10, 1974, p. 8.

Halloran, R. "Japan's Economy in a Deep Slump," *New York Times*, June 12, 1974, pp. 61 and 67.

Japan Times, "Noted Economist Says: Wage Increases Will Help Push Up Prices," June 21, 1974, p. 9.

Japan Times, "Ohira Pledges Efforts to Gain Price Stability," August 3, 1974, p. 1.

South China Morning Post, "Japan Plans Controls to Fight Price Rises," October 8, 1974, p. B2.

Japan Times, "15% 'Limit' to Price Increases Agreed On," October 12, 1974, p. 1.

Editorial. "Priority to Inflation Fight." *Japan Times*, October 23, 1974, p. 12.

Japan Times, "Tight Govt. Budget for 1975 Foreseen," October 29, 1974, p. 9.

Japan Times, "Fukuda Urges Labor to Help End Inflation," December 19, 1974, p. 8.

Japan Times, "Govt. Will Keep Tight Money Policy: Fukuda," January 18, 1975, p. 9.

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Koshiba, K. "Business Pressing for Bigger Cut." *Japan Times*, April 21, 1975, p. 11.

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Japan Times, "Nationwide Consumer Price Increase Hit 9.4% for 1976," April 29, 1977, p. 1.

Editorial. "The New Reflation Program." *Japan Times*, September 8, 1977, p. 16.

Japan Times, "Ohira Pushes for 7% Growth Rate in Debate," January 22, 1978, p. 3.

Japan Times, "Bank of Japan Lowers Official Discount Rate to Record 3.5%," March 16, 1978, p. 1.

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Inflation-Forecast-Based Rules and Indeterminacy: A Puzzle and a Resolution*

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We examine an interesting puzzle in monetary economics between what monetary authorities claim (namely, to be forward looking and preemptive) and the poor stabilization properties routinely reported for forecast-based rules. Our resolution is that central banks should be viewed as following "Calvo-type" inflation-forecast-based (IFB) interest rate rules that depend on a discounted sum of current and future rates of inflation. Such rules might be regarded as both within the legal frameworks and potentially mimicking central bankers' practice. We find that Calvo-type IFB interest rate rules are, first, less prone to indeterminacy than standard rules with a finite forward horizon. Second, in difference form, the indeterminacy problem disappears altogether. Third, optimized forms have good stabilization properties as they become more forward looking, a property that sharply contrasts that of standard IFB rules. Fourth, they appear data coherent when incorporated into a well-known estimated dynamic stochastic general equilibrium (DSGE) model of the euro area.

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1. Introduction

All modern central banks stress the importance of forward-looking policy, and many stress the notion that interest rates should be based on future inflation expectations. Well-known examples include the central banks of Canada and New Zealand, but this is also true by implication of the practice of other central banks. For example, the monetary policy strategy of the European Central Bank (ECB) states that "price stability is to be maintained over the medium term" (ECB 1999, 47), which precisely suggests a forward-looking non-inflationary strategy. The basis for inflation-forecast-based (IFB) rules is that, by anchoring expectations, they improve the credibility and transparency of monetary policy as well as allow policy to be preemptive.

However, such rules have been criticized on various fronts. One concerns the result that typical forward-looking monetary policy rules (such as Taylor-type rules) tend to lead to real indeterminacy (Woodford 2003, chap. 4). This implies that when a shock displaces the economy from its equilibrium, there are an infinite number of possible paths for the real variables leading back to equilibrium. Such "sunspot equilibria" are of interest because sunspot fluctuations—i.e., persistent movements in inflation and output that materialize even in the absence of shocks to preferences or technology—are typically welfare reducing and potentially quite large. Whether policy rules lead to real indeterminacy depends on whether feedback parameters in the policy rules are insufficiently, or indeed overly, aggressive as well as on the length of the forecast horizon itself (e.g., see Levin, Wieland, and Williams 2003, Batini et al. 2006, and the references therein).

Consequently, there would appear to an interesting puzzle in monetary economics between what policymakers routinely claim (namely, to be forward looking and preemptive) and the poor stabilization properties reported for forecast-based rules. Given the importance of aligning central banks' communication strategies with the modeling frameworks underlying much of their technical analysis, this puzzle should not only be taken seriously but indeed would appear to require an urgent resolution. The purpose of this paper is to suggest such a resolution. We propose viewing central banks as following "Calvo-type" IFB interest rate rules that depend on a discounted sum of current and all future rates of inflation. Such rules, it turns out, are less prone to indeterminacy than standard ones with a finite forward horizon, and, if formulated in difference form, the indeterminacy problem disappears altogether. Indeed, we show that optimized Calvo-type rules have good stabilization properties as they become more forward looking, a property that sharply contrasts that of standard IFB rules. Finally, when taken to the data, they appear to behave at least as well as and sometimes better than more standard monetary-policy reaction functions.

Abstracting from such technical characteristics, moreover, the Calvo-type rules we examine might also be regarded as both within the legal framework of and potentially mimicking central bankers' practice. To illustrate, consider the following tension related to the literal use of forecast-based rules. On the one hand, we know that authorities frequently tailor policy and communication strategies to forward-looking outcomes. On the other, we further know that forward-looking policy rules are susceptible to indeterminacy. Moreover, central banks themselves generate expectations for future-dated outcomes but do so in a chronically uncertain environment. Accordingly, we might conjecture that while policymakers will want to incorporate forecasts into their decision strategies, they may be reluctant to treat them commensurate with realized outcomes.

This "chronically uncertain environment" faced by policymakers takes many forms. Consider a few examples. First, macroeconomic time series tend to be actively revised in the quarters following their publication; thus, rule-based policy prescriptions derived from realized data may depart significantly from their real-time counterparts (e.g., Orphanides 2001). It goes without saying that forecasting in such a "noisy" data environment complicates the policy process considerably; authorities might then take recourse to contemporaneous or backward-looking rules, or else persevere with strategies that explicitly incorporate but potentially downplay (i.e., discount) forward-looking information.

Second, and more fundamentally, central banks often employ strong conditioning assumptions in these very forecasts, such as prescribed projections of financial variables, shock processes, external assumptions, and so on. Potentially, therefore, forecasts (particularly medium-term ones) might be considered more a benchmark for scenario analysis and discussion than a specific expected outturn.

In line with this, forecasts are often wrapped around confidence intervals or "fan charts" whose widths are necessarily increasing in the forecast horizon. Moreover, with every new forecast round, data, assumptions, expert judgment, and risks are updated such that forecasts themselves may be heavily revised over time and differ markedly across institutions. Again, in such circumstances, central banks may wish to incorporate these forecasts in their information set but weigh them accordingly. Similarly, one might consider other germane examples based on the various forms of model and judgmental uncertainty (e.g., Onatski and Stock 2002 and Svensson 2005) and its consequences for attenuated or non-attenuated policymaking. Summing up, we might say that while forward-looking policies require forecasts, the very nature of the policy process—i.e., forecast and judgmental revisions, real-time data problems, model uncertainty, etc.—constrains policymakers to treat such information in a manner different from realized outcomes. Our solution—to think of policy as a Calvo-type IFB rule—though simple, is quite powerful: policymakers target future outcomes (such as future inflation rates) in a geometrically discounted manner. We show that this precludes indeterminacy for a number of cases and, indeed, appears data coherent when appended to an estimated dynamic stochastic general equilibrum (DSGE) model.

The paper proceeds as follows. Section 2 sets out a model chosen for its tractability and summarizes the analytical findings of the literature on IFB rules in their standard form. In the analysis, we focus exclusively on "pure" inflation targeting without a feedback on the output gap. We do this for two reasons: First, there are problems associated with measuring the output gap and therefore implementing rules of the "Taylor" type. Second, since simplicity per se is regarded as a positive aspect of monetary rules, it is of interest to study the stabilizing performance of rules that are indeed as simple and transparent as possible. The new contribution to the IFB literature is in sections 3 and 4. Section 3 introduces and analyzes the general properties of a Calvo-type interest rate rule. Section 4 compares such a rule with more conventional j-period-ahead IFB rules and the benchmark of fully optimal monetary policy. Section 5 illustrates the relative empirical performance of our chosen rule when incorporated into the well-known Smets-Wouters DSGE model of the euro area. Section 6 concludes.

2. The Model and Previous Results for IFB Rules

We adopt a standard and tractable New Keynesian model popularized notably by Clarida, Galí, and Gertler (1999) and Woodford (2003):

$$\pi_t = \beta E_t \pi_{t+1} + \lambda mc_t \tag{1}$$

$$mc_t = -(1+\phi)a_t + \sigma c_t + \phi y_t \tag{2}$$

$$c_t = E_t c_{t+1} - \frac{1}{\sigma}(i_t - E_t \pi_{t+1}) \tag{3}$$

$$y_t = c_y c_t + g_y g_t \tag{4}$$

$$a_t = \rho_a a_{t-1} + \epsilon_{a,t} \tag{5}$$

$$g_t = \rho_g g_{t-1} + \epsilon_{g,t}. \tag{6}$$

In (1) and (3), π_t is the inflation rate, β is the private sector's discount factor, $E_t(\cdot)$ is the expectations operator, y_t is output, c_t is consumption, and the slope of the Phillips curve λ can be expressed in terms of the average contract length of Calvo-type price contracts. mc_t given by (2) is the marginal cost, where a_t is a technology shock and ϕ is the Frisch parameter. (1) is derived as a linearized form of staggered price setting about a zero-inflation steady state, and (3) is a linearized Euler equation with i_t the nominal interest rate and σ the risk-aversion parameter. (4) is a linearized aggregate equilibrium relation, where g_t is a government-spending shock and c_y and g_y are consumption and government-spending shares, respectively, in the steady state. According to (5) and (6), shocks follow AR(1) processes. All variables are expressed as deviations from the steady state; π_t and i_t as absolute deviations; and c_t, y_t, and g_t as proportional deviations.

To close the model, we require an interest rate rule. The cornerstone of much of the monetary policy literature is the well-known Taylor (1993) rule, a generalized version of which is

$$\begin{aligned} \rho \in [0,1) : i_t &= \rho i_{t-1} + (1-\rho)\big[\pi_t^* + \theta_\pi E_t\big(\pi_{t+j} - \pi_{t+j}^*\big) \\ &\quad + \theta_y E_t(y_{t+k} - \hat{y}_{t+k})\big] \\ \rho = 1 : i_t &= i_{t-1} + \big[\Theta_\pi E_t\big(\pi_{t+j} - \pi_t^*\big) + \Theta_y E_t(y_{t+k} - \hat{y}_{t+k})\big], \end{aligned} \tag{7}$$

where i_t is the nominal interest rate, π_t is inflation over the interval $[t-1, t]$, $\hat{y}$ is potential output, and y_t is actual output, so that $y - \hat{y}_t$ is the output gap. Variables i_t, $\hat{y}_t$, and y_t are measured in deviation form about a zero-inflation steady state and π_t^* is the inflation target. Integers j, k are the policymaker's forecast horizons, which are a feedback on single-period inflation over the interval $[t+j-1, t+j]$ and a feedback on the output gap over the period $t+k$. Thus, this specification of an interest rate rule accommodates not only outcome-based rules (with $j, k \leq 0$) but also forecast-based ones (with $j, k > 0$). Finally, $\theta_\pi, \theta_y > 0$ and $\Theta_\pi, \Theta_y > 0$ are feedback parameters: the larger the values of these parameters, the faster the pace at which the central bank acts to eliminate the gap between expected inflation and the expected output gap and their target values.

The parameter $\rho \in [0, 1]$ measures the degree of interest rate smoothing. If $\rho = 1$, we have an *integral (or difference) rule* that is equivalent to the interest rate responding to a *price-level target*.[1] For $\rho < 1$, (7) can be written as $\Delta i_t = \frac{1-\rho}{\rho}[\theta_\pi E_t(\pi_{t+j} - \pi^*_{t+j}) + \theta_y E_t(y_{t+k} - \hat{y}_{t+k}) - i_t]$, which is a partial adjustment to a static IFB rule, $i_t = \theta_\pi E_t(\pi_{t+j} - \pi^*_{t+j}) + \theta_y E_t(y_{t+k} - \hat{y}_{t+k})$.

For reasons already discussed, our analysis focuses on standard IFB rules without an output-gap target ($\theta_y = 0$) and with a zero-inflation target $\pi_t^* = 0$.[2] Then, writing $\theta_\pi \equiv \theta$ and $\Theta_\pi \equiv \Theta$, (7) becomes

$$\begin{aligned} i_t &= \rho i_{t-1} + \theta(1-\rho)E_t\pi_{t+j}; \rho \in [0, 1), \theta > 0 \\ &= i_{t-1} + \Theta E_t\pi_{t+j}; \rho = 1, \Theta > 0. \end{aligned} \tag{8}$$

Stability and indeterminacy of a dynamic system are associated with the roots of the system's characteristic equation or,

[1] Unlike its non-integral counterpart, an integral rule responding to inflation does not require observations of the steady-state (natural) rate of interest, about which i_t is expressed, to implement. The merits of price-level versus inflation targeting are examined in Vestin (2006).

[2] Another form of IFB rule found in the literature targets *average* inflation over a specified time horizon, as investigated by Batini and Pearlman (2002). This is represented as $i_t = \rho i_{t-1} + \theta(1-\rho)\frac{E_t \sum_{r=0}^{j} \pi_{t+r}}{1+j}$. As indicated after result 4 below, these rules have roughly the same determinacy properties as a standard IFB rule with half the horizon j.

equivalently, the eigenvalues of its state-space setup. If the number of unstable roots (outside the unit circle) exactly matches the number of nonpredetermined variables, there is a unique solution path. Too few unstable roots then leads to indeterminacy, while too many leads to instability.

The mechanism through which indeterminacy arises can be illustrated in the context of a simplified version of our model. On the demand side, we replace the Keynes-Ramsey condition (3) with an ad hoc IS curve $y_t = -\alpha(i_t - E_t\pi_{t+1})$ and assume $y_t = c_t$ and $\beta = 1$. We also remove the productivity shock a_t. Moreover, suppose that the central bank employs a non-integral rule without interest rate smoothing ($\rho = 0$) so that (7) becomes $i_t = \theta E_t\pi_{t+1}$. Substituting out for y_t and i_t, we arrive at the following process for inflation:

$$E_t(\pi_{t+1}) = \frac{1}{1 - \lambda(\sigma + \phi)\alpha(\theta - 1)}\pi_t. \qquad (9)$$

Consider the case in which private-sector expectations are driven by a nonfundamental-shock process and anticipate that inflation next period will be equal to 1. This will lead to an increase in real interest rates, with a consequent reduction in demand of $\alpha(\theta - 1)$. Given (9), price-setting behavior will thus imply a current inflation rate of $1 - \lambda(\sigma + \phi)\alpha(\theta - 1)$, which we define as π_0.

Now assume that θ is chosen so that $0 < \pi_0 < 1$, which is the case if $1 + 1/(\lambda(\sigma + \phi)\alpha) > \theta > 1$. If we then lead equation (9) forward in time and take expectations, consistency requires that the sequence of successive inflationary expectations is given by 1, $1/\pi_0, 1/\pi_0^2, 1/\pi_0^3, \ldots$. However, these inflation expectations tend to infinity—a solution that clashes with private-sector expectations. Thus, the unique possible solution is $\pi_t = y_t = i_t = 0$ for all $t > 0$. On the other hand, suppose that the central bank is *not aggressive*, and $\theta < 1$. In this case, $\pi_0 > 1$, and hence the sequence of inflationary expectations tends to zero—a solution that fulfills private-sector expectations, making these "self-fulfilling." Now suppose the central bank is *overaggressive* such that $\pi_0 < -1$. This happens when $\theta > 1 + \frac{2}{\lambda(\sigma + \phi)\alpha}$. In this case, the economy experiences cycles of positive and negative inflation, but again the sequence of inflationary expectations tends to zero and fulfills private-sector expectations.

"Self-fulfilling expectations" implies that any initial private-sector expectation leads to an acceptable path for inflation—hence indeterminacy. Furthermore, if these (nonfundamental) shocks to private-sector expectations follow a stochastic process, then sunspot equilibria are generated. These are typically welfare reducing because they induce increased volatility in the system.

So far, research on monetary policy strategy has identified a series of circumstances under which forward-looking optimal and simple one-period-ahead IFB rules might result in multiple equilibria or instability. One of the earliest contributions on indeterminacy under inflation-targeting forward-looking rules is Bernanke and Woodford (1997). Assuming that agents form their expectations rationally, they showed that the equilibrium associated with forward-looking optimal inflation-targeting rules under commitment may not be unique when the central bank targets current (exogenously determined) private-sector forecasts of inflation, either those made explicitly by professional forecasters or those implicit in asset prices. In this sense, their finding squares with the more general one in Sargent and Wallace (1975), who showed that any policy rule responding uniquely to exogenous factors may induce multiple rational-expectations equilibria.

Subsequent work by Svensson and Woodford (2005), again assuming rational expectations and commitment on the side of the central bank, revealed, however, that forward-looking optimal inflation targeting based on endogenously determined forecasts as opposed to exogenous, private-sector forecasts might not necessarily lead to superior results. As their work emphasizes, the purely forward-looking procedure, often assumed in discussions of inflation-forecast targeting, prevents the target variables from depending on past conditions. In other words, the target variables are not "history dependent."[3] This feature makes the rules suboptimal, perhaps seriously so (Currie and Levine 1993), and can lead to indeterminacy of the equilibrium (Woodford 1999).

Perhaps the best-known theoretical result in the literature on IFB rules is that to avoid indeterminacy, the monetary authority must respond aggressively—i.e., with a coefficient above unity,

[3]As we shall see in section 5, this is a property of the optimal-commitment rule.

but not excessively large, to expected inflation in the closed-economy context (see, among others, Clarida, Galí, and Gertler 2000 and, in the small-open-economy context, see De Fiore and Liu 2002). Bullard and Mitra (2002) reaffirmed this result, and in a further paper, Bullard and Mitra (2007) show how monetary policy inertia (a high ρ in (7)) can help alleviate problems of inertia. These same authors follow the pioneering work of Evans and Honkapohja (2001) and examine conditions for the stability of one-period-ahead IFB rules under least-squares learning ("e-stability"). A possible resolution of our "puzzle" is suggested by Evans and Honkapohja (2003) in that forward-looking rules with least-squares learning can lead to stability. However, we address the question of stability and determinacy within a strict rational-expectations framework.

This literature on the determinacy of IFB rules applies to one-period-ahead IFB rules. In a series of papers, Batini and Pearlman (2002), Batini, Levine, and Pearlman (2004), and Batini et al. (2006) extend this literature to any forward horizon for both the closed and open economy. The main results for the former can be summarized as follows.

Result 1. *For an integral rule feeding back on current inflation* ($j = 0$), $\Theta > 0$ *is a necessary and sufficient condition for determinacy.*[4] *For higher feedback horizons* ($j \geq 1$), $\Theta > 0$ *is a necessary but not sufficient condition for stability and determinacy.*

Result 2. *For* j*-period-ahead integral IFB rules,* $j \geq 1$, *there exists a range* $\Theta \in [0, \bar{\Theta}(j)]$ *with* $\bar{\Theta}(j) > 0$ *such that the model is stable and determinate.*

Result 3. *For a non-integral rule feeding back on current inflation* ($j = 0$), $\theta > 1$ *is a necessary and sufficient condition for determinacy. For higher feedback horizons* ($j \geq 1$), $\theta > 1$ *is a necessary but not sufficient condition for stability and determinacy.*

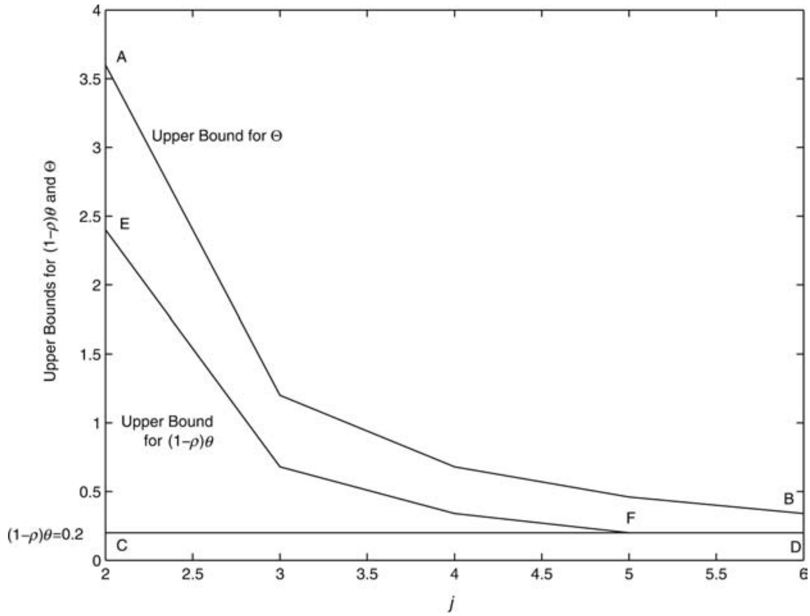
[4]Note that determinacy implies saddle-path stability; however, a model can be stable with non-explosive behavior but without a unique solution (i.e., indeterminate).

Result 4. *For j-period-ahead non-integral IFB rules, $j \geq 1$, there exists some lead J such that for $j > J$, there is indeterminacy for all values of θ.*[5] *J is given by*

$$J = \frac{1}{1-\rho} + \frac{(1-\beta)\sigma}{\lambda(\sigma+\phi)}. \tag{10}$$

To get a feel for these results, we now provide numerical values for threshold values $\bar{\theta}$ for non-integral rules and $\bar{\Theta}$ for integral rules. In figure 1, based on table 1, parameter estimates are taken from Batini et al. (2006).[6] For non-integral rules, we set $\rho = 0.8$.

Figure 1. Critical Upper Bounds for $(1-\rho)\theta$ and Θ



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[5]Strictly, there are some mild conditions on the parameters that a plausible calibration easily satisfies for this result to hold—see Batini and Pearlman (2002). For the average-inflation rule of the type set out in the previous footnote, the corresponding lead $\hat{J}$ is given by $\hat{J} = 2J - 1$.

[6]Parameter values are $\lambda = 0.27$, $\beta = 0.99$, $\sigma = 3.91$, $\phi = 2.16$, $\rho_a = \rho_g = 0.9$, $sd(\epsilon_g) = 2.75$, and $sd(\epsilon_a) = 0.59$, found using Bayesian methods and U.S. data.

Table 1. Critical Upper Bounds for $\bar{\theta}(j)$ and $\bar{\Theta}(j)$

Threshold	ρ	$j=1$	$j=2$	$j=3$	$j=4$	$j=5$	$j=6$
$\bar{\theta}(j)$	0.8	102	12.0	3.4	1.70	1.00	Indeterminacy
$\bar{\Theta}(j)$	1.0	23	3.6	1.2	0.68	0.46	0.34

These numerical results corroborate the analytical results summarized above.[7] The indeterminacy problem becomes more acute as the horizon j increases, imposing a tighter constraint on the range of IFB rules available. For non-integral rules with $\rho = 0.8$, the maximum horizon J is just over five quarters. In accordance with result 2, for integral rules, as j increases, there is always some feedback coefficient on expected inflation $0 < \Theta < \bar{\Theta}$ such that the IFB rule yields stability and determinacy. For non-integral rules, the area of determinacy in $(j, (1-\rho)\theta)$ space is EFC. For integral rules, the corresponding space in (j, Θ) space is ABDC.[8]

3. Calvo-Type Interest Rate Rules

We now turn to the main focus of this paper, which is an alternative way of thinking about IFB rules, referred to in the introduction as *Calvo-type interest rate rules.*[9] To formulate this, first define the discounted sum of future expected inflation rates as

$$\Theta_t = (1-\varphi)E_t(\pi_t + \varphi\pi_{t+1} + \varphi^2\pi_{t+2} + \cdots); \varphi \in (0,1). \tag{11}$$

[7]In fact, qualitatively similar results are found in a more developed New Keynesian model with consumption habit and price indexing in Batini et al. (2006).

[8]Further insight into these results can be provided by writing the expected value of future inflation approximately as $E_t\pi_{t+j} = (\lambda^{max})^j\pi_t$, where λ^{max} is the largest stable eigenvalue of the system under control. Then as j increases, $(\lambda^{max})^j$ decreases, so that the feedback effect becomes negligible, and the system exhibits indeterminacy similar to the $\theta < 1$ type.

[9]We use this terminology since they have the same structure as Calvo-type price or wage contracts (Calvo 1983). One can think of the rule as a feedback from expected future inflation that continues in any one period with probability φ and is switched off with probability $1-\varphi$. The probability of the rule lasting for just j periods is then $(1-\varphi)\varphi^j$, and the mean-lead horizon is therefore $(1-\varphi)\sum_{j=1}^{\infty} j\varphi^j = \frac{\varphi}{1-\varphi}$.

Then

$$\varphi E_t \Theta_{t+1} - \Theta_t = -(1-\varphi)\pi_t. \tag{12}$$

With this definition, a rule of the form

$$\begin{aligned} i_t &= \rho i_{t-1} + \theta(1-\rho)\Theta_t; \rho \in [0,1), \theta > 0 \\ &= i_{t-1} + \Xi\Theta_t; \rho = 1, \Xi > 0 \end{aligned} \tag{13}$$

emerges that describes feedback on forward-looking inflation with mean-lead horizon $\frac{\varphi}{1-\varphi}$. Thus, with $\varphi = 0.5$, e.g., we have a Calvo-type rule that compares with (7) with a horizon $j = 1$.[10]

Consider first *non-integral* rules. With a Calvo-type rule—writing (1), (3), (12), and (13) in matrix form—the characteristic equation of the system can be shown to be

$$\begin{aligned} (1-\varphi z)(z-\rho)((\beta z - 1)(z-1) - \frac{\lambda(\sigma+\phi)}{\sigma} z) \\ + \theta(1-\varphi)(1-\rho)\frac{\lambda(\sigma+\phi)}{\sigma} z = 0, \end{aligned} \tag{14}$$

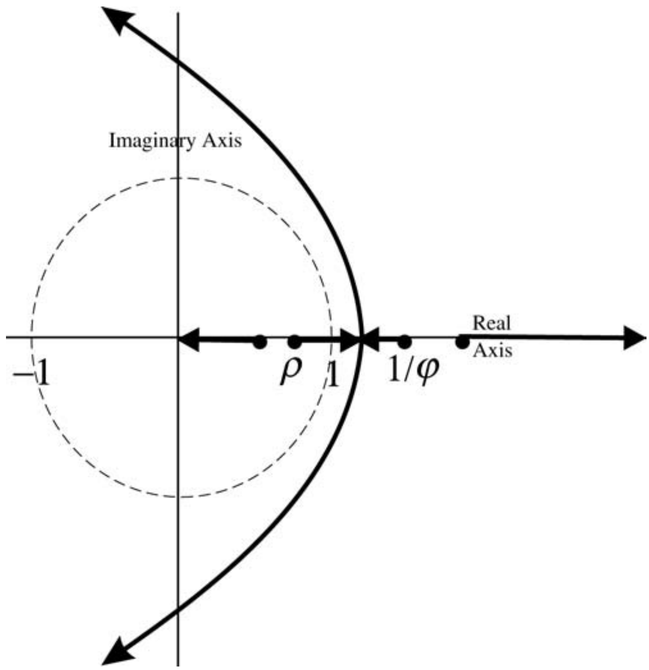
where z is the forward operator (i.e., $zx_t \equiv x_{t+1}$). Noting that the system (1), (3), and (13) has only one lag term, the condition for stability and indeterminacy of the system is that exactly one root of (14) must lie within the unit circle. Accordingly, we investigate (14) using the *root-locus method.* This is a standard method for analyzing the stability of dynamic linear systems found in the engineering literature. (See Evans 1954 and Aoki 1981 for an early

[10]It is of interest to note that for $\rho \in [0,1)$, this rule can also be expressed as

$$i_t = (1+\varphi\rho)^{-1}[\rho i_{t-1} + \varphi E_t i_{t+1} + \theta(1-\rho)\pi_t].$$

Whether the rule is expressed in this way or as (13), it is evident from the model (1)–(6) that current variables and one-step-ahead forecasts are sufficient statistics for the decisions of private agents. Under least-squares learning of the rule, the private sector does not need to look further than one period, but with fully rational expectations, assumed in this paper, the rule expressed in terms of $E_t i_{t+1}$ or (13) must be solved forward, getting us back to the original infinite-horizon formulation (11).

Figure 2. Position of Roots as θ Changes from 0 to ∞



application to economics.)[11] It was used for the first time to study the indeterminacy of forward-looking interest rate rules by Batini and Pearlman (2002). The method provides an elegant way of locating the position in the complex plane of all the roots of the characteristic equation as one of the parameters changes. In our application, the parameter in question is the feedback parameter from future inflation, θ.

The root-locus diagram for (14) as θ changes is shown in figure 2, which depicts the complex plane, and is a generic shape for all parameter values of the system. The root locus starts out at the roots of (14) for $\theta = 0$; these roots are denoted on the diagram by $\bullet$. Note that one root, $z = 1/\varphi$, is outside the unit circle, while another, $z = \rho$, is inside the unit circle, and it is easy to show that $(\beta z - 1)(z - 1) - \frac{\lambda(\sigma+\phi)}{\sigma} z = 0$ has one root outside and one root inside

[11] See appendix 1 for a brief guide to the root-locus method.

the unit circle. The arrows then show how the four roots change as θ changes. Note in particular that the smallest root has a branch from it leading to $z = 0$, while the largest has a branch leading to $z = \infty$. Of the other two roots, one of them has a branch passing through $z = 1$, and where their branches meet, they both branch into the complex plane and head to infinity at an angle of 60° asymptotically to the real line.[12]

There are several things to note about this root-locus diagram. First, when $\theta = 1$, then $z = 1$ as well; this is immediate from (14).[13] Second, the diagram therefore implies that the system has a single stable root for all values of $\theta > 1$, no matter what the values are of the other parameters. However, this apparently general result needs some explanation and, indeed, some slight qualification. We summarize the main results as follows and provide a proof in appendix 2.

Result 5. *A sufficient condition for the system (1)–(3) with the Calvo-type interest rate rule (13) to be determinate for all* $\theta > 1$ *is that* $\rho > \varphi$.

Estimated interest rate rules (including our estimates in section 5) suggest substantial smoothing, with typically $\rho > 0.95$. The condition in this last result is therefore that $\varphi < 0.95$, or, in other words, the mean lead must be less than nineteen quarters. A final observation is that as the interest rate smoothing increases, at the limit where we have an integral rule, result 5 *always* holds. Thus the result for integral (price-level) rules is an immediate corollary of result 5.

Result 6. *The system (1)–(3) with the Calvo-type integral (i.e., price-level) interest rate rule is determinate for all* $\Xi > 0$.

The proof follows once one has replaced $(1 - \rho)\theta$ with Ξ, so that the same argument follows as for the proof of the previous result (with $z = 1$ when $\Xi = 0$) and, in addition, $1 > \varphi$.

---

[12] The root-locus diagram would look qualitatively the same if $1/\varphi$ were the largest real root, or ρ the smallest real root, for $\theta = 0$.

[13] Note that the Taylor principle—that interest rate should react by more than one-to-one to expected or current or past inflation—means that $\theta > 1$ for non-integral rules.

Thus, according to result 6, the indeterminacy problem *disappears altogether* (in the context of our simple model) if the authorities target a weighted average of present and future price levels with geometrically declining weights.

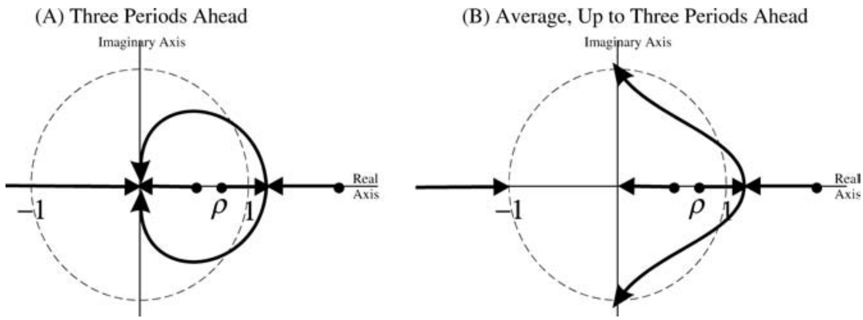
The Calvo-type IFB rule (13) is not completely forward looking, as it includes a reaction to current inflation with weight unity. Do the improved determinacy properties of this rule compared with standard IFB rules crucially depend on the presence of current inflation? We can see this is not the case because average-inflation rules, which also react to current inflation over a finite time horizon j referred to in footnotes 2 and 3, have similar indeterminacy properties to single-period IFB rules with horizon $\frac{j+1}{2}$ if j is odd and $\frac{j}{2}$ if j is even. Further suppose that the Calvo-type rule involves only forward-looking inflation, so that (12) contains the term $E_t\pi_{t+1}$ instead of π_t, so that the characteristic equation (14) becomes

$$(1-\varphi z)(z-\rho)((\beta z-1)(z-1)-\frac{\lambda(\sigma+\phi)}{\sigma}z) + \theta(1-\varphi)(1-\rho)\frac{\lambda(\sigma+\phi)}{\sigma}z^2=0. \tag{15}$$

Using the root-locus technique, one can show that, provided $\rho > \varphi$, there is indeterminacy only for values of θ beyond that value at which $z=-1$. Thus the critical value of θ for indeterminacy is that which satisfies (15) at $z=-1$. It is easy to see that this critical value is given by $f(\rho)(1+\varphi)/(1-\varphi)$, where $f(\rho)$ is a function of ρ (and the other parameters), which is an increasing function of φ. Thus for the Calvo-type rule, as φ (and therefore the expected horizon) increases, the proneness to indeterminacy actually *falls*. Furthermore, the function $f(\rho)\to\infty$ as $\rho\to 1$, so with Calvo-type *integral* rules that are purely forward looking, the critical value for θ above which there is indeterminacy becomes infinite and result 6 holds.

We end this section by noting the contrast between result 5 and result 3, which can be illustrated by the root-locus diagrams of figure 3 for IFB rules. These diagrams depict the cases for interest rates depending on either single-period inflation, three periods ahead (panel A), or average inflation over the current period and up

Figure 3. Position of Roots for (A) Single-Period-Inflation Forward-Looking Rules and (B) Average-Inflation Forward-Looking Rules as θ Changes from 0 to ∞



to three periods ahead (panel B). Both diagrams demonstrate that there may be a range of $\theta > 1$ for which there is determinacy (exactly one stable root), but for values of θ that are too large, there is indeterminacy.

4. Optimal Monetary Policy

4.1 Utility-Based Welfare

In the simple model of this paper, a quadratic approximation to the utility of the household that underlies the model takes the form

$$\Omega_0 = E_0\left[\frac{1}{2}\sum_{t=0}^{\infty}\beta^t\left[(y_t - \hat{y}_t)^2 + w_\pi\pi_t^2 + w_i i_t^2\right]\right], \tag{16}$$

where $\hat{y}_t$ is potential output achieved when prices are flexible ($mc_t = 0$ in (1)) and

$$w_\pi = \frac{\zeta}{\lambda(\sigma+\phi)} = \frac{\zeta\xi}{(1-\xi)(1-\beta\xi)(\sigma+\phi)}. \tag{17}$$

In (17), $1-\xi$ is probability of a price optimization for each firm, σ is the risk-aversion parameter, $1+\phi$ is the elasticity of disutility with respect to hours worked, and ζ is the elasticity of substitution of differentiated goods making up aggregate output.[14] For estimated or calibrated parameter values (see footnote 6 above) reported in Batini et al. (2006), this gives $w_\pi = 1.826$.[15]

Although there is no cost-push ("markup") shock in our model, the existence of a penalty on the variability of the interest rate, driven by concerns for the zero-lower-bound constraint, leads to an inflation/output-gap trade-off. Whereas without such a constraint optimal policy simply sets the interest rate to keep both inflation and the output gap at zero, with the constraint this is not possible and a nontrivial policy problem emerges.

4.2 Optimal Policy with and without Commitment

We first compute the optimal policies where the policymaker can commit and the optimal discretionary policy where no commitment mechanism is in place.[16] To obtain the weight on interest rate variance, w_i, we first compute the optimal-commitment rule with the interest rate responding only to current inflation (see below). We impose an approximate *zero lower bound* on the nominal interest rate by experimenting with w_i so that it is sufficiently high so as to ensure $i_t > 0$ with almost unit probability of 0.99 percent; i.e., (assuming a normal distribution), $sd(i_t) < \frac{i}{2.33}$, where $i = (\frac{1}{\beta} - 1) \times 100$ is the natural rate of interest. A weight $w_i = 0.5$ was necessary to achieve this condition.

4.3 Optimized IFB Rules

We now turn to optimized IFB rules. The general form of the rule that covers both integral and non-integral rules is given by

$$i_t = \rho i_{t-1} + \Xi E_t \pi_{t+j}; \rho \in [0,1], \Xi, j \geq 0. \tag{18}$$

[14] See Woodford (2003, chap. 6).

[15] Based on an annual inflation rate, this is equivalent to $w_\pi = \frac{1.826}{16} = 0.11$, which is at the lower end of commonly used weights.

[16] Full details of the solution procedures for optimal commitment and discretion are given in Levine, McAdam, and Pearlman (2007). For the former we show that, in general, optimal commitment can be implemented as a rule that feeds back on both current and past predetermined state variables and is therefore history dependent. See also Woodford (2003, chap. 8) for ways in which optimal commitment can be implemented as a rule.

The corresponding Calvo-type rules are given by

$$i_t = \rho i_{t-1} + \Xi E_t \Theta_t; \rho \in [0, 1], \Xi, j \geq 0. \tag{19}$$

Given the estimated variance-covariance matrix of the white-noise disturbances, an optimal combination (Ξ, ρ) can be found for each rule defined by the time horizon $j \geq 0$.

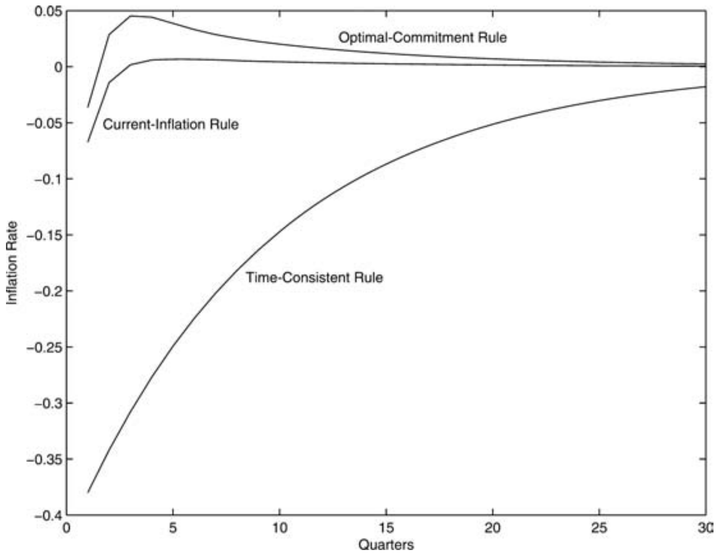
4.4 Numerical Results

We first focus on the optimal-commitment rule, the optimal-discretionary (time-consistent) rule, and an optimized current-inflation rule. Results for the three types of rules are summarized in table 2. Figures 4–11 compare the responses under the three rules following an unanticipated productivity shock ($a_0 = 1$) and an unanticipated government-spending shock ($g_0 = 1$). The output and inflation equivalent welfare differences compared with the optimal-commitment policy are computed as follows. Suppose the welfare

Table 2. Comparision of Welfare-Based Optimal Rules and Optimized IFB Rules

Rule	ρ	Ξ	**Loss Function**	y_e	π_e
Minimal Feedback on π_t	1	0.001	49.01	0.97	0.72
IFB(0)	1	2.035	2.509	0.11	0.08
IFB(1)	1	12.00	2.676	0.12	0.09
IFB(2)	1	3.570	4.574	0.23	0.17
IFB(3)	1	1.216	32.02	0.78	0.57
IFB(4)	1	0.675	208.1	2.03	1.50
Calvo IFB($\varphi = 0.5$)	1	2.203	2.602	0.12	0.09
Calvo IFB($\varphi = 0.67$)	1	2.351	2.636	0.12	0.09
Calvo IFB($\varphi = 0.75$)	1	2.444	2.653	0.12	0.09
Calvo IFB($\varphi = 0.875$)	1	2.616	2.678	0.13	0.09
Calvo IFB($\varphi = 0.917$)	1	2.683	2.689	0.13	0.09
Optimal Commitment	n.a.	n.a	1.896	0.00	0.00
Optimal Discretion	n.a.	n.a	53.55	1.02	0.75

Figure 4. Inflation following Shock $a_0 = 1$



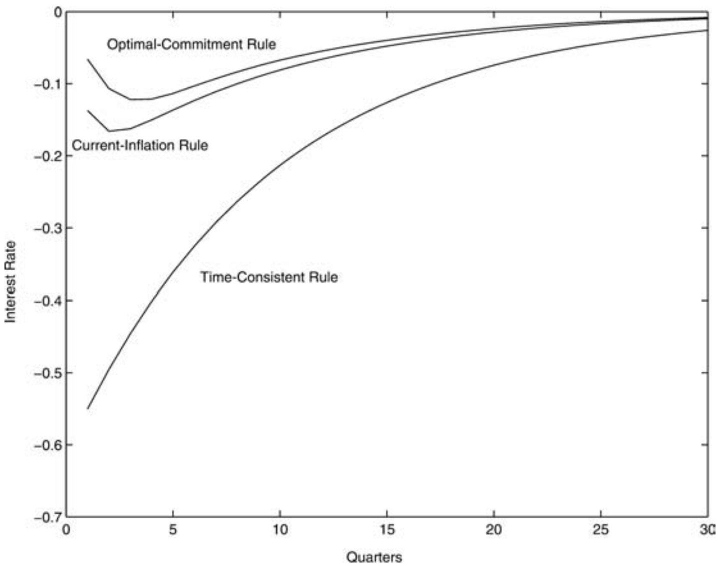
loss difference is X. This is equivalent to a permanent output gap of y_e if $\frac{1}{2(1-\beta)} y_e^2 = X$ and to a permanent inflationary bias of π_e if $\frac{1}{2(1-\beta)} w_\pi \pi_e^2 = X$; i.e.,

$$y_e = \sqrt{2(1-\beta)X} \tag{20}$$

$$\pi_e = \sqrt{\frac{2(1-\beta)X}{w_\pi}}. \tag{21}$$

Comparing the three types of rules, there are two notable results. First, Clarida, Galí, and Gertler (1999) stress the existence of stabilization gains from commitment in New Keynesian models: we show in our simple model that these are substantial, amounting to a permanent output equivalent of 1.02 percent or an inflationary bias of 0.75 percent per quarter, or 3 percent per year. The source of this time-inconsistency problem is from pricing and consumption behavior together. Following a shock that diverts the economy from its steady state, given expectations of inflation, the opportunist

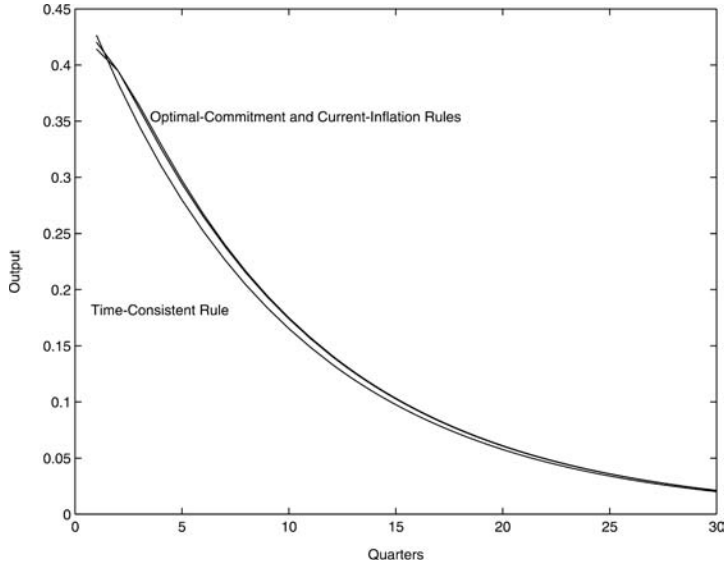
Figure 5. Interest Rate following Shock $a_0 = 1$



policymaker can increase or decrease output by reducing or increasing the interest rate, which increases or decreases inflation. Consider the case where the economy is below its steady-state level of output. A reduction in the interest rate then causes consumption demand to rise. Firms that are locked into price contracts respond to an increase in demand by increasing output and increasing the price according to their indexing rule. Those that can reoptimize only increase their price. These changes are for given inflationary expectations and illustrate the incentive to inflate when the output gap increases. In a noncommitment equilibrium, however, the incentive is anticipated, and the result is greater inflation variability as compared with the commitment case. This contrast between the commitment and discretionary cases is seen clearly in the figures.

The second notable result concerns the optimized current-inflation rule. We find that most of the gains from commitment (in fact, over 80 percent) can be achieved by this very simple optimized rule without an output-gap feedback, and the cost of simplicity is

Figure 6. Output following Shock $a_0 = 1$



only 0.11 percent of output, or an inflationary bias of 0.08 per quarter, or 0.32 percent per year. We find the optimized rule over parameters ρ and Ξ in (18) gives $\rho = 1$ so that the best current-inflation rule is of the integral type. In the figures, we see how the optimized current-inflation rule closely mimics the optimal-commitment rule.

Turning now to IFB rules, we compute the optimized standard rules with future horizon $j = 0, 1, 2, 3, 4$, denoted by IFB(j), and compare these with Calvo-type rules with probability of survival $\varphi = 0.5, 0.67, 0.75, 0.875$, and 0.917 corresponding to an average future horizon of $\frac{\varphi}{1-\varphi} = 1, 2, 4, 7$, and 11 quarters, respectively. Our results first confirm a finding of Batini et al. (2006): that the stabilization performance of standard optimized IFBj rules deteriorates sharply as the horizon j increases. Our new result that follows from the stability analysis of Calvo-type rules and the absence of an "indeterminacy constraint" is that *this sharp deterioration is not a feature of Calvo-type optimized IFB rules.* Even optimized rules with an expected future horizon of three years perform almost as well as the current-inflation rule. Again, we find that integral rules perform the best.

Figure 7. Output Gap following Shock $a_0 = 1$

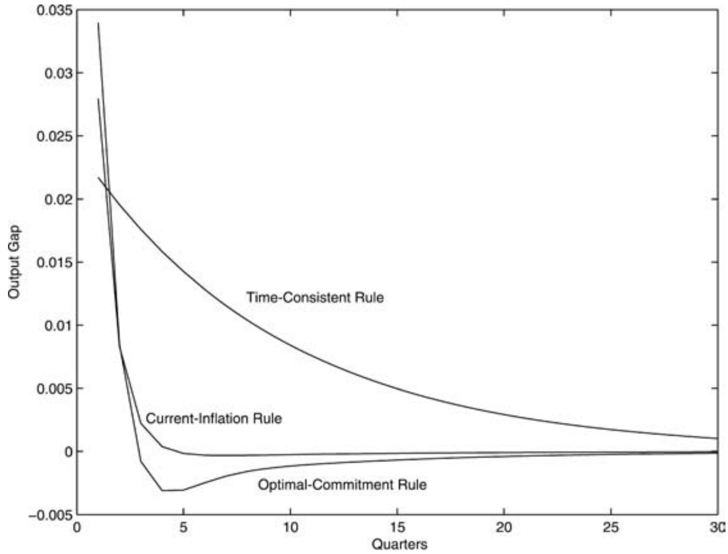


Figure 8. Inflation following Shock $g_0 = 1$

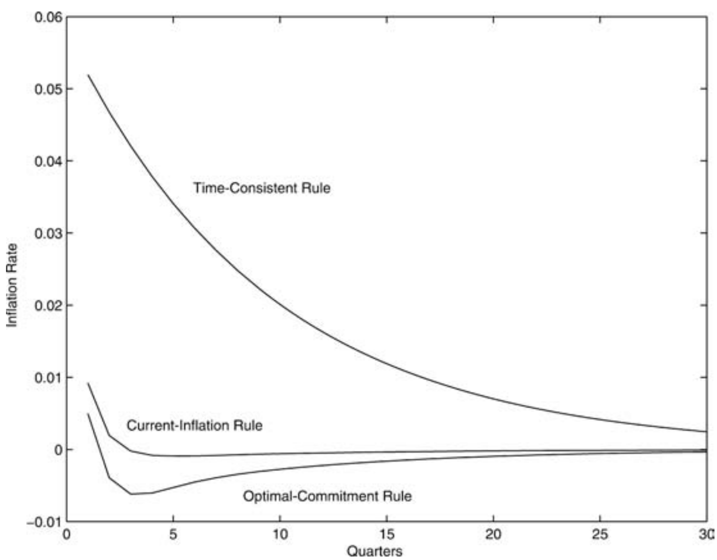


Figure 9. Interest Rate following Shock $g_0 = 1$

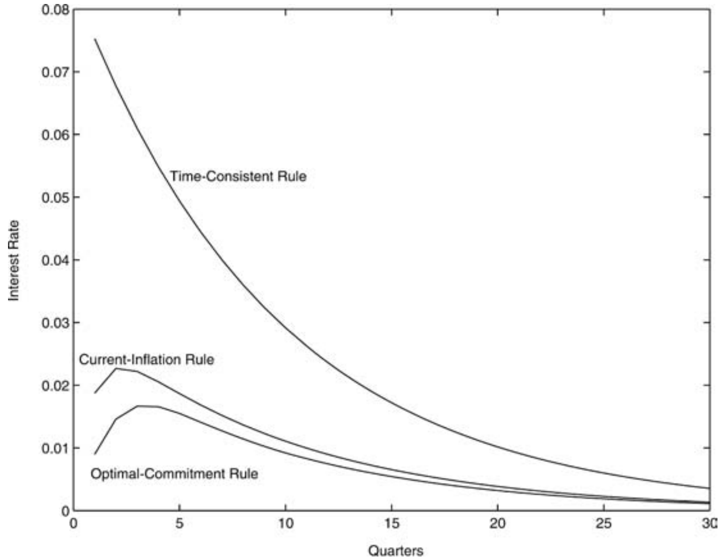


Figure 10. Output following Shock $g_0 = 1$

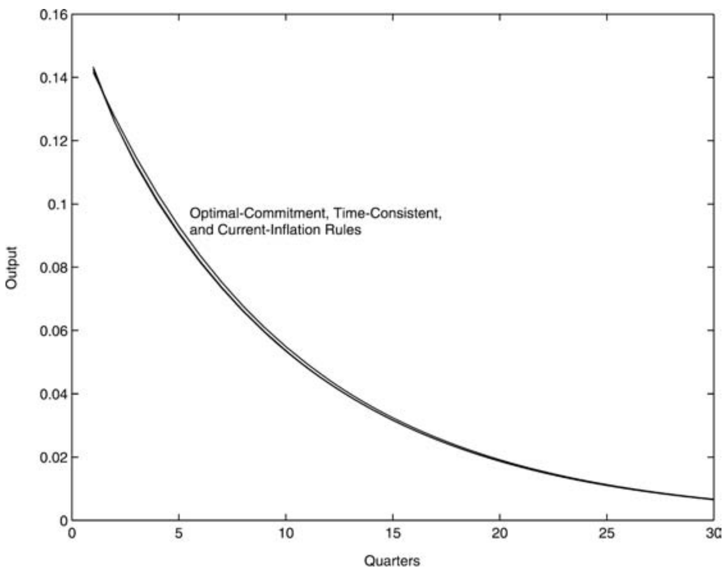
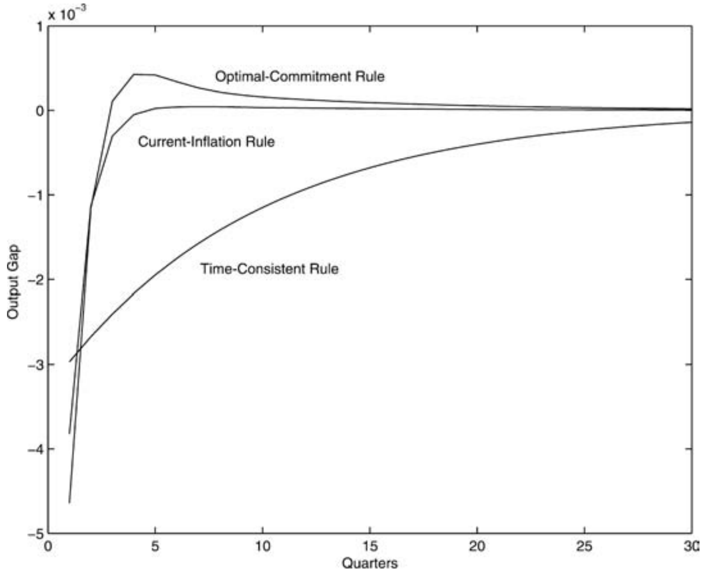


Figure 11. Output Gap following Shock $g_0 = 1$



5. Calvo-Type Interest Rate Rules: A DSGE Model Illustration

In this section, for illustrative purposes, we implement the aforementioned Calvo-type interest rate rules in a benchmark model of the euro area—namely, that of Smets and Wouters (2003) (SW henceforth). A brief description of the model is provided in appendix 3.

5.1 Monetary-Policy Reaction Functions

In line with the empirical approach to monetary rules in SW, we modify the previous monetary-policy reaction functions for the standard and Calvo-type IFB rules as, respectively,

$$\begin{aligned} i_t &= \rho i_{t-1} + (1-\rho)[\bar{\pi}_t + \theta_\pi E_t(\pi_{t+j} - \bar{\pi}_{t+j}) + \theta_y \tilde{y}_t] \\ &\quad + \theta_{\Delta\pi}(\pi_t - \pi_{t-1}) + \theta_{\Delta y}(\tilde{y}_t - \tilde{y}_{t-1}) \\ i_t &= \rho i_{t-1} + (1-\rho)[\bar{\pi}_t + \theta_\pi \Theta_t(\varphi) + \theta_y \tilde{y}_t] \\ &\quad + \theta_{\Delta\pi}(\pi_t - \pi_{t-1}) + \theta_{\Delta y}(\tilde{y}_t - \tilde{y}_{t-1}), \end{aligned} \tag{22}$$

where $\tilde{y}_t$ is the output gap. Then in the SW model $j = -1$ and the interest rate feeds back on lagged inflation. To incorporate this rule as a special case, we also modify (11) to become

$$\Theta_t = (1-\varphi)E_t\big[\pi_{t-1} - \pi^*_{t-1} + \varphi\big(\pi_t - \pi^*_t\big) + \varphi^2\big(\pi_{t+1} - \pi^*_{t+1}\big) + \cdots\big];\ \varphi \in (0,1), \tag{23}$$

so (12) now becomes

$$\varphi E_t\Theta_{t+1} - \Theta_t = -(1-\varphi)\big(\pi_{t-1} - \pi^*_{t-1}\big). \tag{24}$$

This modified Calvo-type rule reduces to the past-inflation-rate rule in SW as a special case by putting $\varphi = 0$. The mean-lead horizon is now given by $\frac{1}{1-\varphi} - 2$.[17]

5.2 Results

Thus, we reestimate by Bayesian methods the SW model with the policy rule replaced by (22) and the model supplemented with (23), where, to repeat, the special case of $\varphi = 0$ retrieves the default, backward-looking SW policy rule.[18] Table 3 reports the parameters of the policy reaction function for each model variant, from IFB(−1) to IFB(4) to the Calvo case.[19] As standard, two sets of parameter results are presented: (i) the estimated posterior mode of the parameters, which is obtained by directly maximizing the log of the posterior distribution with respect to the parameters (and a standard error based on the corresponding Hessian) and (ii) the 5th and 95th percentile of the posterior distribution of the parameters

[17] It is straightforward to show that the results of section 3 still hold with this modification.

[18] We are grateful to Gregory De Walque and Raf Wouters for providing the SW model in Dynare code.

[19] Results for the other parameters (as well as the Dynare files to replicate our results) are available on request from the authors. Notably, the full set of parameter values appeared very well identified and stable across the model variants.

Table 3. Comparison of Calvo-Type and Standard IFB Rules

Rule		**IFB(−1)**	**IFB(0)**	**IFB(1)**	**IFB(4)**	**Calvo-Type IFB**
ρ	Mode	0.969	0.969	0.965	0.951	0.958
		(0.012)	(0.013)	(0.015)	(0.025)	(0.021)
	Mean	0.965	0.967	0.958	0.940	0.951
		[0.943:0.984]	[0.951:0.986]	[0.932:0.983]	[0.891:0.977]	[0.918:0.982]
θ_π	Mode	1.700	1.701	1.700	1.700	1.702
		(0.099)	(0.100)	(0.100)	(0.099)	(0.099)
	Mean	1.697	1.698	1.700	1.703	1.707
		[1.535:1.853]	[1.531:1.860]	[1.542:1.873]	[1.541:1.867]	[1.539:1.868]
θ_y	Mode	0.121	0.121	0.117	0.111	0.120
		(0.045)	(0.045)	(0.046)	(0.049)	(0.045)
	Mean	0.117	0.123	0.109	0.107	0.120
		[0.045:0.186]	[0.057:0.193]	[0.034:0.178]	[0.028:0.171]	[0.053:0.186]
$\theta_{\Delta\pi}$	Mode	0.146	0.146	0.111	0.118	0.121
		(0.052)	(0.052)	(0.049)	(0.049)	(0.049)
	Mean	0.155	0.110	0.116	0.119	0.126
		[0.070:0.239]	[0.034:0.195]	[0.034:0.200]	[0.041:0.196]	[0.048:0.212]
$\theta_{\Delta y}$	Mode	0.154	0.154	0.152	0.147	0.151
		(0.023)	(0.023)	(0.022)	(0.022)	(0.021)
	Mean	0.152	0.152	0.146	0.139	0.146
		[0.120:0.191]	[0.115:0.187]	[0.112:0.183]	[0.099:0.178]	[0.109:0.181]
φ	Mode	–	–	–	–	0.8398
		–	–	–	–	(0.103)
	Mean	–	–	–	–	0.797
		–	–	–	–	[0.646:0.956]
Prob.		0.289	0.096	0.158	0.224	0.234

Note: IFBj rule: $i_t = \rho i_{t-1} + (1-\rho)[\pi_t^* + \theta_\pi E_t(\pi_{t+j} - \pi_{t+j}^*) + \theta_y \tilde{y}_t] + \theta_{\Delta\pi}(\pi_t - \pi_{t-1}) + \theta_{\Delta y}(\tilde{y}_t - \tilde{y}_{t-1})$. Calvo IFB ($\varphi$): $i_t = \rho i_{t-1} + (1-\rho)[\pi_t^* + \theta_\pi \Theta_t(\varphi) + \theta_y \tilde{y}_t] + \theta_{\Delta\pi}(\pi_t - \pi_{t-1}) + \theta_{\Delta y}(\tilde{y}_t - \tilde{y}_{t-1})$, where $\varphi E_t \Theta_{t+1} - \Theta_t = -(1-\varphi)(\pi_{-t} - \pi_{t-1}^*)$. Hessian standard errors are in parentheses and 5th and 95th percentiles are in squared brackets. Log marginal likelihood of IFB(−1) = −298.65.

obtained through the Metropolis-Hastings sampling algorithm (using 100,000 draws from the posterior, three parallel chains, and an average acceptance rate of around 0.25) for the various model variants. The models are estimated using the Dynare software (Juillard 2004). Note that, in reestimation, we used identical priors to those used in SW. Moreover, for the additional parameter, φ, we assumed a beta distribution with a prior mean of 0.8 (corresponding to a mean-lead horizon of three quarters), with a standard error of 0.1.

Turning to the results, we see that the Calvo rule yields a φ value centered at an implied mean-lead horizon of three to four quarters in the policy rule. As shown in the last row of table 3, which reports the model odds, this rule beats all contemporaneous and forecast-based rules in marginal-likelihood terms without leading to any deterioration in the parameter values. Comparing the likelihood values of the Calvo rule with the backward-looking rule, IFB(–1), there is a very close data coherence. Indeed, in terms of Bayesian odds ratio (0.234/0.289 = 0.81), we effectively could not discriminate between these two types of rules.[20] Summing up, one might say that while by no means conclusive, these results do suggest that a Calvo-type rule is competitive with more conventional monetary policy rules.

6. Conclusions

The large literature on IFB rules now strongly suggests that rules that target future inflation with a specified time horizon are prone to indeterminacy and have poor stabilization properties. This raises

[20] As discussed in Geweke (1999), the Bayesian approach to estimation allows a formal comparison of different models based on their marginal likelihoods. The marginal likelihood of model M_i is given by

$$p(Y \mid M_i) = \int_{\Xi} p(\xi \mid M_i) p(Y \mid \xi, M_i) d\xi,$$

where $p(\xi \mid M_i)$ is the prior density for model M_i and $p(Y \mid \xi, M_i)$ is the data density for model M_i given the parameter vector ξ and the data vector Y. Then the posterior odds ratio is given by

$$PO_{ij} \equiv \frac{p(M_i|Y)}{p(M_j|Y)} = \frac{p(Y|M_i)p(M_i)}{p(Y|M_j)p(M_j)} = \frac{p(Y|M_i)}{p(Y|M_j)},$$

assuming equal prior model probabilities ($p(M_i) = p(M_j)$). The posterior model probabilities are reported in table 3.

an interesting puzzle of why many central banks insist on forward-looking inflation targets. Part of the answer lies in the fact that these rules assume commitment to a low long-run inflation rate (in fact, zero inflation in our setup). For central banks, forward-looking rules with a low inflation target signals commitment to this low long-run inflation rate. But can IFB rules with a long forward lead be implemented without negative consequences? Our paper proposes a resolution of this puzzle by suggesting that the policy process of central banks may in fact be best modeled in the form of a Calvo-type rule that targets a discounted infinite sum of future expected inflation.

Our main findings are, first, that Calvo-type IFB interest rate rules are less prone to indeterminacy than standard ones with a finite forward horizon. Second, for such rules in integral (i.e., difference) form, the indeterminacy problem disappears altogether. In this case, the Calvo rule takes the form of a weighted-average future price-level target with geometrically declining weights. Third, as a consequence of these results, optimized Calvo-type rules have good stabilization properties as they become more forward looking, which sharply contrasts with the substantial deterioration in the corresponding performance of standard IFB rules. Fourth, in terms of data coherence in the context of the SW model, a Calvo-type rule with a mean forward horizon of just less than one year is perfectly competitive with more conventional monetary policy rules.

A number of possible directions for future research are suggested by this study. First, Calvo-type reaction functions can be estimated directly using generalized method of moments (GMM) estimation methods and can be compared with more standard Taylor-type and IFB rules. Second, in DSGE modeling in general (open economy, closed economy, interacting economies, etc.), it is commonplace to compare the performance of optimized Taylor-type rules with their optimal counterparts. Calvo-type rules could be added to this exercise. Finally, the design of robust rules using, e.g., the Bayesian estimated posterior distribution as in Batini et al. (2006) could be extended to rules of this form.[21]

[21]See Levine et al. (2007).

Appendix 1. A Topological Guide to the Root-Locus Technique

Here we present a brief guide on how to use the root-locus technique. We start with some standard rules as provided in control-theory textbooks and then apply them to the specific example of the paper.

The idea is to track the roots of the polynomial equation $f(z)+\theta g(z) = 0$ as θ moves from 0 to ∞. Clearly for $\theta = 0$, the roots are those of $f(z) = 0$, whereas when $\theta \to \infty$, the roots are those of $g(z) = 0$. The root locus then connects the first set of roots to the second set by a series of lines and curves. We shall assume without loss of generality that the coefficient of the highest power of f is negative and that of g is positive. Since the roots of a polynomial may be complex, the root locus is plotted in the complex plane.

There are a number of different ways to state the standard rules that underlie the technique. One popular way (see Evans 1954) of sketching the root locus by hand involves just six steps:

(i) a. Define $n(f)$ = number of zeros of $f(z)$, $n(g)$ = number of zeros of $g(z)$. For our case, $n(f) = 4, n(g) = 1$.

b. Loci start at the zeros of $f(z)$ and end at the zeros of $g(z)$ and at ∞ if $n(f) > n(g)$.

(ii) Number of loci must be equal to $\max(n(f), n(g)) = 4$, in our case.

(iii) A point on the real axis is on the root locus if the number of zeros of f and g on the real axis to its left is odd.

(iv) Loci ending at ∞ do so at angles to the positive real axis given by $2k\pi/(n(f) - n(g))$, where the integer k ranges from 0 to $(n(f) - n(g)) - 1$. In our case, these angles are 0, $2\pi/3$, $4\pi/3$.

(v) If all coefficients of f and g are real, then the root locus is symmetric about the real axis.

(vi) Loci leave the real axis where $\partial\theta/\partial z = 0$.

Appendix 2. Proof of Result 5

We prove this result in two steps. First, we need to show that the branch point into the complex plane near $z = 1$ is to the right of $z = 1$. Second, we have to show that the branches of the root locus do not cross the unit circle twice (otherwise, there are too many stable roots, and hence indeterminacy, over a certain range of values of θ greater than 1).

STEP 1. The branch point is to the right of $z = 1$, provided that the root locus passes through this point from left to right as θ increases. But this means that we require $\partial z/\partial\theta > 0$ at $z = \theta = 1$. By implicit differentiation of (14), we find that

$$\left[(1-\beta)(1-\rho)(1-\varphi) + \frac{\lambda(\sigma+\phi)}{\sigma}(1-2\varphi+\rho\varphi)\right]\left.\frac{\partial z}{\partial\theta}\right|_{\theta=1}$$
$$= (1-\rho)(1-\varphi).$$

It is easy to see that a sufficient condition for $\partial z/\partial\theta|_{\theta=1} > 0$ is $\rho > \varphi$.

STEP 2. We now investigate those points on the root loci that lie on the unit circle. These are, of course, characterized by $z = e^{i\phi} = cos\phi + isin\phi$. To solve for ϕ, the easiest approach is to substitute $z = e^{i\phi}$ directly and then multiply (14) through by $e^{-i\phi}$. Then the imaginary part of this expression is independent of θ and can be written as

$$\begin{aligned}\varphi\beta sin3\phi - [\beta(1+\varphi\rho) + \varphi(1+\beta+\lambda(\sigma+\phi)/\sigma))]sin2\phi \\ + [(1+\varphi\rho)(1+\beta+\lambda(\sigma+\phi)/\sigma) \\ + \varphi + \rho\beta - \rho]sin\phi = 0. \qquad (25)\end{aligned}$$

Using the substitutions $sin2\phi = 2sin\phi\cos\phi$, $sin3\phi = (4cos^2\phi - 1)$ $sin\phi$, it is clear that one solution to (25) is $sin\phi = 0$, which corresponds to $\phi = 0$ ($z = 1$) and $\phi = \pi$ ($z = -1$, which is technically a solution when $\theta < 0$). It follows that the other solutions are given by

$$\begin{aligned} 4\varphi\beta cos^2\phi - 2[\beta(1+\varphi\rho) + \varphi(1+\beta+\lambda(\sigma+\phi)/\sigma))]cos\phi \\ + (1+\varphi\rho)(1+\beta+\lambda(\sigma+\phi)/\sigma) \\ + \varphi + \rho\beta - \rho - \varphi\beta = 0. \end{aligned}$$

Provided that $\rho > \varphi$, it is easy to show that the coefficient of $cos\phi$ is more than twice that of $cos^2\phi$; it follows that at least one of the solutions to $cos\phi$ is greater than 1. But this means that there is no more than one real solution for ϕ, so that there cannot be a double crossing of the unit circle for $\rho > \varphi$.

Appendix 3. The Smets-Wouters Model

The Smets-Wouters (SW) model is an extended version of the standard New Keynesian DSGE closed-economy model with sticky prices and wages. The model features three types of agents: households, firms, and the monetary policy authority. Households maximize a utility function with two arguments (goods and leisure) over an infinite horizon. Consumption appears in the utility function relative to a time-varying external habit-formation variable. Labor is differentiated over households, so that there is some monopoly power over wages, which results in an explicit wage equation and allows for the introduction of sticky nominal Calvo-type wage contracts. Households also rent capital services to firms and decide how much to accumulate given certain capital-adjustment costs. Firms produce differentiated goods, decide on labor and capital inputs, and set Calvo-type price contracts. Wage and price setting is augmented by the assumption that those prices and wages that cannot be freely set are partially indexed to past inflation. Prices are therefore set as a function of current and expected real marginal cost but are also influenced by past inflation. Real marginal cost depends on wages and the rental rate of capital. The short-term nominal interest rate is the instrument of monetary policy. The stochastic behavior of the model is driven by ten exogenous shocks—five shocks arising from technology and preferences, three cost-push shocks, and two monetary policy shocks.

Consistent with the DSGE setup, potential output is defined as the level of output that would prevail under flexible prices and wages in the absence of cost-push shocks.

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Modern Forecasting Models in Action: Improving Macroeconomic Analyses at Central Banks*

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There are many indications that formal methods are not used to their full potential by central banks today. In this paper, using data from Sweden, we demonstrate how BVAR and DSGE models can be used to shed light on questions that policymakers deal with in practice. We compare the forecast performance of BVAR and DSGE models with the Riksbank's official, more subjective forecasts, both in terms of actual forecasts and root mean-squared errors. We also discuss how to combine model- and judgment-based forecasts, and show that the combined forecast performs well out of sample. In addition, we show the advantages of structural analysis and use the models for interpreting the recent development of the inflation rate through historical decompositions. Last, we discuss the monetary transmission mechanism in the models by comparing impulse-response functions.

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1. Introduction

Over the last two decades, there has been a growing interest among macroeconomic researchers in developing new tools for quantitative monetary policy analysis. In this paper, we examine how the new generation of formal models can be used in the policy process at an inflation-targeting central bank. We compare official forecasts published by Sveriges Riksbank (the central bank of Sweden) with forecasts from two structural models—a dynamic stochastic general equilibrium (DSGE) model and an identified vector autoregression (VAR) estimated with Bayesian methods. We also discuss how the formal models can be used for storytelling and for the analysis of alternative policy scenarios, which is an important matter for central banks.

It is rather unusual that formal models are contrasted to official central bank forecasts. Such a comparison is especially interesting given that the latter also include judgments and "extra-model" information, which are very hard to capture within a formal setting. Hence, this can give an assessment of how useful expert knowledge is and shed light on the "role of subjective forecasting" raised by Sims (2002). An evaluation of official forecasts and formal model forecasts has, to our knowledge, only been conducted for the Board of Governors of the Federal Reserve System (see Altig, Carlstrom, and Lansing 1995 and Sims 2002).[1] We will supplement this analysis by evaluating official and model forecasts for a small open economy (Sweden).

Smets and Wouters (2004) have shown that modern *closed-economy* DSGEs (with various nominal and real frictions) have forecasting properties well in line with more empirically oriented models such as standard and Bayesian VARs (BVARs). However, this paper evaluates forecasts from DSGE and BVAR models that include *open-economy* aspects and compares these with judgmental forecasts. Given the increased complexity of open-economy models, and the different monetary policy transmission mechanism where exchange rate movements are of importance, our exercise adds an

[1] Before this paper was completed, Edge, Kiley, and Laforte (2006) compared the forecasting performance of the Federal Reserve with a closed-economy DSGE as well as with a theoretical reduced-form model.

extra element to the analysis in, e.g., Sims (2002) and Smets and Wouters (2004). In section 2, we show the actual inflation and interest rate forecasts from the various setups as well as the root mean-squared errors (RMSEs) of the forecasts. In addition, we explicitly examine two episodes where official and DSGE forecasts diverge to look further into the role of subjective forecasting.

Both the DSGE and BVAR models have now been used within the policy process at Sveriges Riksbank between one and two years. However, this does not provide us with enough observations to carry out an extensive forecast evaluation in genuine real time. Since the DSGE model uses data on as many as fifteen macroeconomic variables, we have therefore estimated the formal models using a revised data set, which makes a much longer evaluation period of the various forecasting methods possible. We discuss the role of real-time versus revised data in detail in section 2 below.

In addition, we demonstrate how formal models can shed light on practical policy questions. In section 3.1, we let the two models interpret the underlying reasons for the recent economic development, using a historical decomposition of the forecast errors in each of the two models. We also show that a VAR model can be used to clarify what has happened in the economy, as long as we are willing to impose some structure to identify the underlying shocks. However, when one is interested in predictions conditioned upon alternative policy scenarios, an idea about how monetary policy is designed and how it affects the economy is required. By comparing impulse-response functions in section 3.2, we show that the DSGE model, using structure from economic theory, provides a much more reasonable transmission mechanism of monetary policy than the BVAR model. This is a necessary requirement for producing conditional forecasts that are meaningful from the perspective of a central banker.

Finally, section 4 summarizes our views on the advantages of formal methods and the reasons why such methods have not been more influential at central banks.

2. Forecasting Performance

This section provides an evaluation of the inflation, interest rate, and GDP forecasts by the Riksbank, a small-open-economy DSGE model, and a Bayesian VAR model. Both actual forecasts and root mean-squared errors are examined for the period 1999:Q1–2005:Q4.[2] To characterize the forecasting advantage of the Riksbank's judgmental forecasts, we also analyze two specific episodes where the different forecasting approaches diverge for CPI inflation and where we, a priori, expect the sector experts to have an informational advantage. Since the Riksbank's forecasts have until recently been intended to be conditioned on the assumption of a constant short-term interest rate, we also look at the interest rate forecasts from the formal models.[3] These are compared with implicit-interest-rate forecasts calculated from (market-based) forward interest rates. Finally, we also look at the GDP forecasts from the different models, but since no genuine real-time data set has been compiled for all fifteen variables in the DSGE model, these results should be interpreted with some caution.[4]

[2]Official inflation forecasts from the Riksbank cannot be obtained on a quarterly basis before 1999:Q1. Moreover, the DSGE model needs a sufficient number of observations after the transition from a fixed exchange rate to the inflation-targeting regime in 1993. Since GDP forecasts are not available from the Riksbank before 2000:Q1, their precision is evaluated between 2000:Q1 and 2005:Q4.

[3]Between October 2005 and February 2007, the Riksbank produced forecasts conditioned upon implicit forward rates, instead of the constant-interest-rate assumption. Since February 2007, the Riksbank has published its preferred path for the future repo rate (i.e., the short-term interest rate controlled by the Riksbank).

[4]The absence of a real-time data set implies that the formal models possibly have an information advantage relative to the official forecasts since the BVAR and DSGE forecasts are based on ex post data on GDP. (Preliminary GDP data are available with a delay of around one quarter, but they are subsequently revised.) This is not a problem with the other variables in the models, since data on prices, interest rates, and exchange rates are available on a monthly basis and are not revised. On the other hand, the official forecasts have a small information advantage in some quarters, when the *Inflation Report* has been published toward the end of the quarter and, thus, can be based on data on prices, interest rates, and exchange rates from the early part of the same quarter. Since most of the variables in the BVAR are available in real time, and we use a Litterman (i.e., random-walk) prior for the lag polynomial, there are good reasons to believe that the BVAR results with the exception of GDP growth are not very sensitive to our decision to use revised data. The DSGE model, which is estimated on a data set that contains several additional real quantities, is probably more sensitive to the use of revised data.

The Riksbank publishes official forecasts in its quarterly *Inflation Report* publication. The forecasts are not the outcome of a single formal model but rather the result of a complex procedure with input both from many different kinds of models and judgments from sector experts and the Riksbank's executive board.[5]

The BVAR model contains quarterly data for the following seven variables: trade-weighted measures of foreign GDP growth in logs (y_f), CPI inflation (π_f) and a short-term interest rate (i_f), the corresponding domestic variables (y, π, and i), and the level of the real exchange rate defined as $q = 100(s + p_f - p)$, where p_f and p are the foreign and domestic CPI levels (in logs) and s is the (log) trade-weighted nominal exchange rate. More details on the BVAR are provided in appendix 1.

The DSGE model is an extension of the closed-economy models developed by Altig et al. (2003) and Christiano, Eichenbaum, and Evans (2005) to the small-open-economy setting in a way similar to that of Smets and Wouters (2002). Households consume and invest in baskets consisting of domestically produced goods and imported goods. We allow the imported goods to enter both aggregate consumption and aggregate investment. By including nominal rigidities in the importing and exporting sectors, we allow for short-run incomplete exchange rate pass-through to both import and export prices. The foreign economy is exogenously given by a VAR for foreign inflation, output, and the interest rate. The DSGE model is estimated with Bayesian methods using data on the following fifteen variables: GDP deflator inflation, real wage, consumption, investment, real exchange rate, short-run interest rate (repo rate), hours, GDP, exports, imports, CPI inflation, investment deflator inflation, foreign (i.e., trade-weighted) output, foreign inflation, and foreign interest rate. The DSGE model is identical to the one developed and estimated by Adolfson et al. (forthcoming), and a more

[5]The forecast process during the evaluation period was of a recursive and iterative nature, where the foreign and financial variables entered first. Given these forecasts, the Swedish real variables were predicted. Typically, the GDP forecast was an aggregation of the components of the GDP identity. The labor-market variables entered in a third step, where the productivity and unit-labor-cost variables were determined. Finally, predictions of CPI and core inflation—mainly based on forecasts of import prices, unit labor cost, and the output gap—ended the first forecast round.

detailed description of the model, priors used in the estimation, and full-sample estimation results can be found in that paper.

The DSGE model contains more variables than the BVAR, but this does not imply that the DSGE model necessarily has an advantage in terms of the forecasting performance, since the inclusion of more variables in the BVAR also implies that a considerably larger number of parameters needs to be estimated. Adolfson et al. (forthcoming) consider a BVAR with the same set of variables as in the DSGE model, and the forecasting performance of this BVAR is not very different from the BVAR used in this paper.[6]

One difficulty when it comes to comparing official inflation forecasts from the Riksbank with model forecasts (or forecasts made by other institutions) is that the Riksbank's forecasts have until recently been intended to be conditioned on the assumption that the short-term interest rate (more specifically, the Riksbank's instrument, the repo rate) remains constant throughout the forecasting period. However, it is not unreasonable to assume that the official forecast actually lies closer to an unconditional forecast, given its subjective nature. In practice, it is extremely difficult to ensure that the judgmental forecast has been conditioned on a constant-interest-rate path rather than on some more likely path.[7] Beyond the forecast horizon, the implicit assumption also seems to have been that the interest rate gradually returns to a level determined by some interest rate equation. Even so, we see from figure 1 (shown on the next set of facing pages) that the interest rate level (annualized average of daily repo-rate observations within each quarter) has been rather stable during the sample period, which is the reason why a

[6]If anything, the introduction of additional variables leads to a reduction in forecasting performance. In particular, this appears to be the case for the nominal interest rate.

[7]The *Inflation Reports* from March and June 2005 contain discussions of the problems with constant-interest-rate forecasts, as well as two forecasts of inflation and GDP growth that are conditional on either a constant interest rate or the implied forward rate. Although there was a considerable difference between the forward rate and the constant interest rate on these occasions (a gradual increase to about 150 basis points at the two-year horizon), the forecasts for GDP and inflation were not very different. This suggests that the official constant-interest-rate forecasts were in fact close to unconditional projections. See also Adolfson et al. (2005) for difficulties with constant-interest-rate forecasts in a model-based environment.

conceivable constant-interest-rate assumption might not have been that unfavorable. Moreover, the impulse-response functions from interest rate changes to inflation are typically fairly small, with a substantial time delay, at least in the BVAR model (see figure 6 in section 3.2), which implies that a constant-interest-rate assumption at long horizons should have relatively small effects on the inflation forecast. We therefore believe that, taken together, our forecast comparisons are valid and our conclusions will not be significantly affected by the alleged constant-interest-rate assumption underlying the official forecasts.

2.1 Inflation Forecasts

Figure 1a presents the outcome of CPI inflation in Sweden for 1998–2005 (bold line) together with the Riksbank's official forecasts (first row), the forecasts from the DSGE model (second row), and the BVAR model (third row) for 1999:Q1–2005:Q4.[8] The visual impression is that the official forecasts and the model forecasts have somewhat different properties. The Riksbank's forecasts appear to be rather "conservative"; most often they predict a very smooth development of inflation, and the changes in the forecast paths are relatively small between quarters. The BVAR model, on the other hand, seems to view the inflation process as more persistent, since forecast errors have a larger influence on subsequent forecasts.

The top panel in figure 2 shows the root mean-squared errors (RMSEs) for different forecast horizons (one to eight quarters ahead) of the yearly CPI-inflation forecasts. The DSGE and the official forecasts have about the same precision for inflation forecasts made up to a year ahead; however, at somewhat longer horizons (five to eight quarters ahead), the forecasts from the DSGE model have a better accuracy than the official inflation forecasts. The BVAR forecasts, on the other hand, perform very well up to six quarters ahead but are beaten by the DSGE's inflation forecasts at longer horizons. It should also be noted that all three forecasting methods are

[8]The data in figure 1 refer to yearly inflation rates ($p_t - p_{t-4}$), just like the inflation series published in the *Inflation Report*. Since the DSGE and BVAR models are specified in terms of quarterly rates of change ($p_t - p_{t-1}$), the model forecasts are summed up to fourth differences.

Figure 1a. Sequential Forecasts of Yearly CPI Inflation, 1999:Q1–2005:Q4, from the Riksbank (First Row), the DSGE Model (Second Row), and the BVAR Model (Third Row)

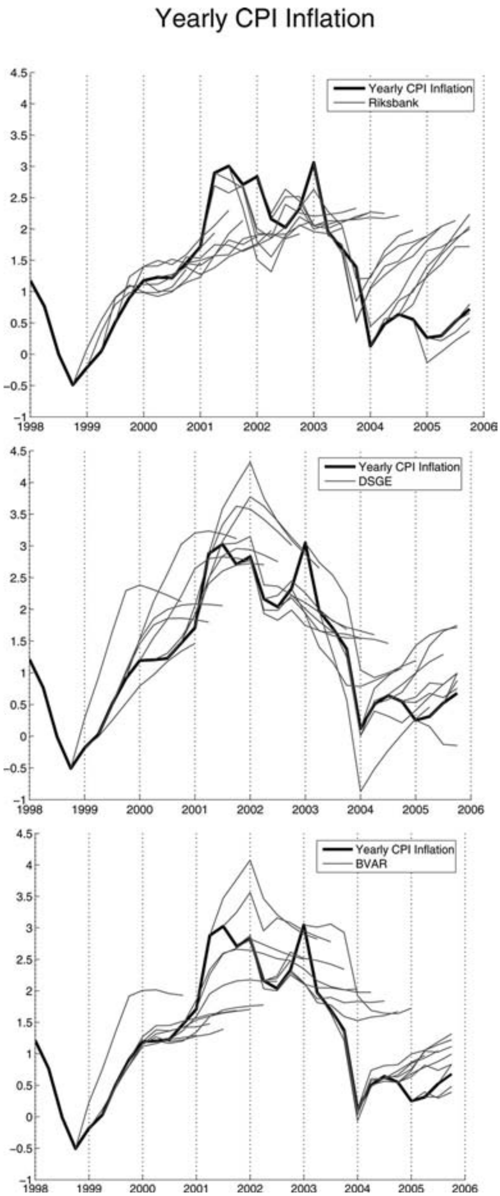


Figure 1b. Sequential Forecasts of Repo Rate, 1999:Q1–2005:Q4, from the Riksbank (First Row), the DSGE Model (Second Row), and the BVAR Model (Third Row)

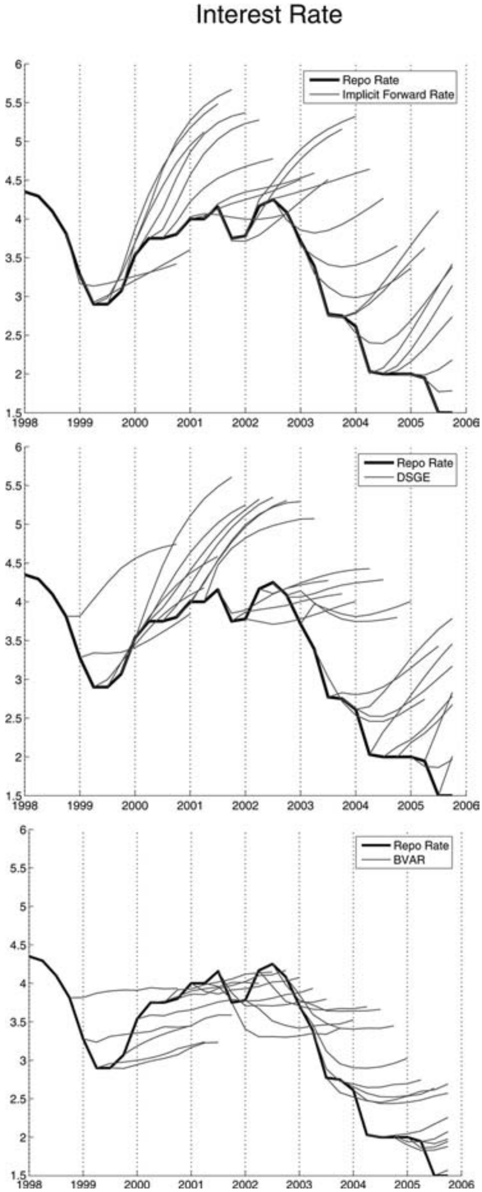


Figure 1c. Sequential Forecasts of Yearly GDP Growth, 1999:Q1–2005:Q4, from the Riksbank (First Row), the DSGE Model (Second Row), and the BVAR Model (Third Row)

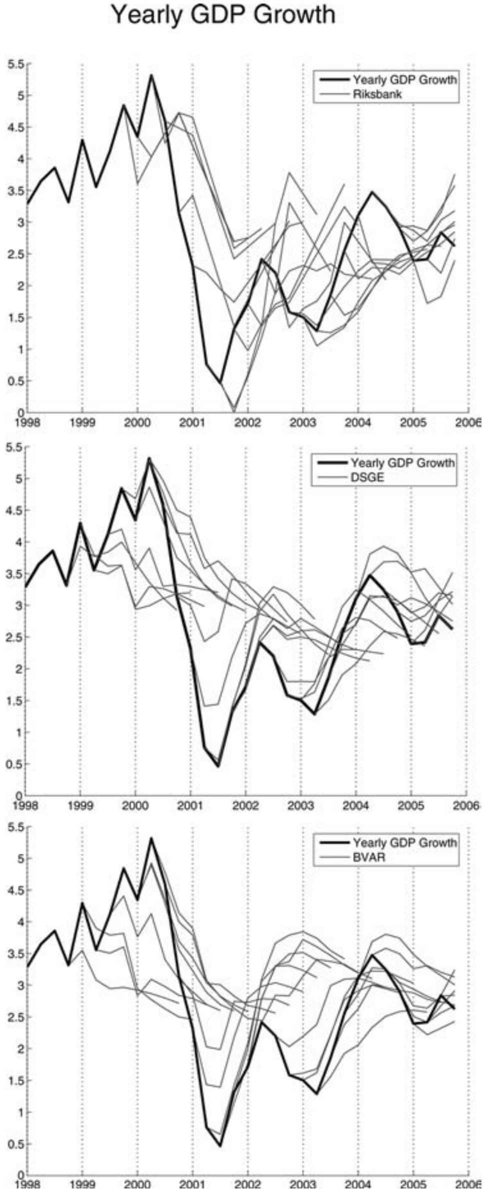
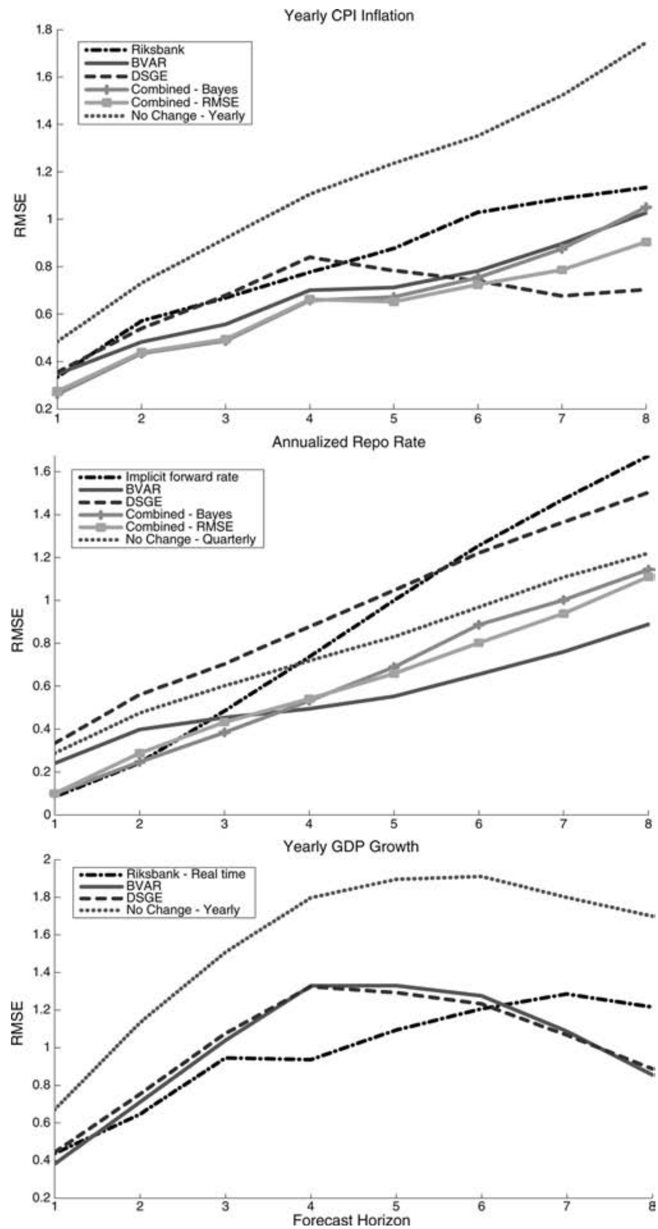


Figure 2. Root Mean-Squared Error (RMSE) of Yearly CPI-Inflation Forecasts (Top) and Annualized Interest Rate Forecasts (Middle) 1999:Q1–2005:Q4, and Yearly GDP-Growth Forecasts (Bottom) 2000:Q1–2005:Q4



better than a naive forecast of constant yearly inflation (the No Change—Yearly line). We take the similar precision in the different inflation forecasts as evidence in favor of all three forecasting approaches. It is encouraging that the model forecasts are performing as well as the official forecasts also at shorter horizons. We will, however, return to the specific advantages and the role of subjective forecasting below.

Although Smets and Wouters (2004) have shown that the forecasting performance of closed-economy DSGE models compares quite favorably to more empirically oriented models such as vector autoregressive models, it is not evident that our DSGE model will have similar properties, given that the open-economy dimensions add complexity to the model. In particular, it is well known that uncovered interest rate parity (UIP) is rejected empirically, which may deteriorate the forecasting performance of an open-economy DSGE model. The UIP condition in the DSGE model has therefore been modified to allow for a negative correlation between the risk premium and expected exchange rate changes; see Adolfson et al. (forthcoming) for further details. Figure 2 reveals that our DSGE model has a forecasting performance for CPI inflation that is remarkably good. The DSGE model makes smaller forecast errors for inflation six to eight quarters ahead than both the BVAR and the official Riksbank forecasts. Part of this can be attributed to the modified UIP condition (see Adolfson et al. forthcoming), but we also believe that the fact that we are considering a stable regime with a known and fixed inflation target makes the theoretical structure imposed by the DSGE model particularly useful.

2.2 Interest Rate Forecasts

Figure 1b shows outcomes of the repo rate along with either the expected interest rate paths derived from forward interest rates (first row), following the method described by Svensson (1995), or along with the repo-rate forecasts from the DSGE model (second row) and the BVAR model (third row). Throughout our sample, forward interest rates have systematically overestimated the future interest rate level. In principle, this could be due to (possibly time-varying) term and risk premia; i.e., forward rates do not provide direct estimates of expectations. But, in practice, we believe it to be more likely that the market (along with the Riksbank and the DSGE model) on average overestimated the inflation pressure and the need for higher nominal interest rates during this period. Turning to the interest rate forecasts from the formal models, we see that future interest rates have been systematically overestimated also in this case, even if the BVAR forecasts have been more moderate than the interest rate forecasts from the DSGE model. The main reason for the relatively large forecast errors of the DSGE model for the repo rate is that the model has tended to overestimate both the inflation pressure and the GDP-growth prospects during the period, which can be seen in figures 1a and 1c. This indicates that it is not the estimated policy rule that does not fit the data: had we used the true inflation and output outcomes when projecting with the rule, the forecast errors for the repo rate would have been greatly reduced.

In the middle panel of figure 2, we compare the RMSEs for the various repo-rate forecasts. It can be seen that forward interest rates, a naive constant-interest-rate forecast, and the BVAR model all have about the same precision for forecasts three to four quarters ahead, whereas the DSGE has somewhat worse accuracy. Naturally, this raises some questions about how monetary policy is described in the DSGE model, although the above reasoning suggests that the policy rule itself may not be the key problem. Forward rates have a better precision one to two quarters ahead, while the BVAR model makes much better forecasts for longer horizons than the other approaches. The fact that the BVAR model provides a more realistic picture than forward rates at longer horizons, which even seem to have a lower predictive power than the constant-interest-rate assumption, may be interpreted as arguments against inflation forecasts conditioned on forward rates. Nevertheless, this is an assumption that the Bank of England recently adopted, and Norges Bank and Sveriges Riksbank recently abandoned. It should be emphasized, however, that our results have been obtained from a sample where the short-term interest rate has been unusually low and stable in a historical context. Thus, it is possible that forward rates are more informative in periods when interest rates change more.

2.3 GDP Forecasts

Figure 1c shows the forecasts for the yearly GDP growth from the Riksbank, the DSGE, and the BVAR models against actual yearly GDP growth (last data vintage in all cases). Given that the Riksbank's forecasts are made in real time, whereas the formal models are estimated using a revised data set for GDP, a direct comparison between them is hard to make. The pattern between the DSGE and BVAR forecasts is relatively similar, however. This also shows up in the root mean-squared errors for the forecasts in figure 2. The models' accuracy in terms of predicting GDP growth is almost the same in this sample (2000:Q1–2005:Q4).

To increase the comparability of the forecasting performances, we compute the RMSEs for the Riksbank's GDP forecasts in real-time data, since the formal models would have an informational advantage if all these GDP forecasts were evaluated on the revised data set. The models appear to perform reasonably well also against the judgmental forecasts of the Riksbank (see figure 2), although it should be emphasized that such a comparison should be interpreted with caution.

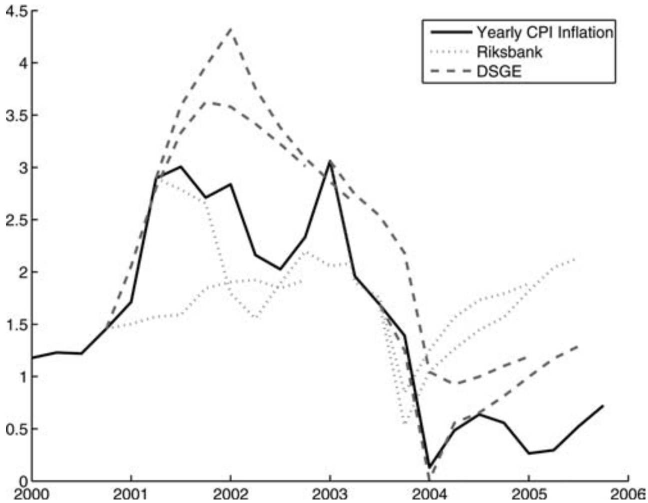
2.4 The Role of Subjective Forecasting

To obtain more information about the properties of the various approaches to forecasting, it is interesting to take a closer look at some specific episodes where the pure model forecast and the official Riksbank forecast differ. This provides us with useful information about whether sector experts can provide a better understanding of recent influences on inflation (which might only be temporary). In figure 3, we compare the Riksbank's official inflation forecasts with the corresponding DSGE forecasts on four different occasions (i.e., figure 3 contains a subset of the information in figure 1).[9]

The first episode concerns forecasts made immediately before and after the sudden increase in inflation in 2001:Q2. One important factor behind this increase was a rise in food prices due to the mad-cow and foot-and-mouth diseases, although the inflation rate had

[9] For ease of exposition, we only focus on the Riksbank and DSGE forecasts here.

Figure 3. Riksbank and DSGE Forecasts of Yearly CPI Inflation Made on Four Specific Occasions: 2000:Q4, 2001:Q2, 2003:Q1, and 2003:Q3



started to increase already in 1999. In 2000:Q4, the Riksbank's official forecast implied a slowly increasing inflation rate over the next two years. The DSGE model suggested a much stronger increase in inflation. This may reflect that the model forecast attributed a larger weight to recent increases in inflation, while the subjective procedure, leading up to the Riksbank's official forecast, underestimated the persistence in changes in inflation. However, once the shock had become apparent, the subjective approach proved to be very useful. At that time, 2001:Q2, the sector experts expected the food price increase to involve a persistent shock to the price level but with small further effects on the yearly inflation rate, and the Riksbank's forecasts at 2001:Q2 were more in line with the actual outcome for the next few quarters. The DSGE model, on the other hand, treated the food price shock as any other inflation shock and overestimated its effects on inflation during both 2001 and 2002.

The second episode concerns inflation forecasts made in 2003:Q1. Cold, dry weather had brought about extreme increases in electricity prices during the winter of 2002/03. This was a temporary shock to the price level, but it had persistent effects on yearly inflation, which became unusually low when electricity prices declined during the spring and summer. The Riksbank's official forecasts described the decline in inflation extremely well for the first two quarters but underestimated the effects on the longer horizons, possibly because it was difficult to separate the effects of changes in energy prices (the oil price also fluctuated heavily) from the downward pressure on inflation from other forces in the economy (e.g., increases in productivity). In contrast, the DSGE model underestimated the drop in inflation at first, but once the decline had started, it more correctly predicted that inflation would be very low for the next one to two years (cf. the forecasts from 2003:Q3 in figure 3).

These episodes nicely illustrate how formal statistical models and judgments by sector experts can complement each other. Subjective forecasts may sometimes be too myopic and pay too little attention to systematic inflation dynamics related to the business cycle or other historically important regularities. Model forecasts, on the other hand, cannot take sufficient account of specific unusual but observable events. At the same time, judgments from sector experts based on their detailed knowledge about the economy can be extremely useful—in particular, when unusual shocks have hit the economy.

2.5 Combined Forecasts

The previous subsections have contrasted model-based and judgment-based macroeconomic forecasts, and have made it clear that judgments from sector experts can be useful in the short run, especially when unusual disturbances to the economy occur. Being equipped with several different forecasts, our natural question is, what can be gained from combining them into a single overall forecast? Combining a set of purely model-based forecasts is rather straightforward, especially within the Bayesian framework, where the weights are given by posterior model probabilities (Draper 1995). When at least one of the forecasting models cannot be represented by a probability model for the observed data, we need to resort to other solutions.

Winkler (1981) proposes an alternative Bayesian approach that may be used to combine model-based and judgment-based forecasts.

We will form weights separately for each variable and forecast horizon. Winkler's procedure is described in detail in appendix 2. The method assumes that the forecast errors from the different forecasts at a specific time period follow a multivariate normal distribution with zero mean and covariance matrix Σ. It is further assumed that the forecast errors are independent over time. The optimal weights on the individual forecasts can then be shown to be a simple function of Σ^{-1}. This means that the weights depend not only on the relative precision of the forecasts but also on the correlation between forecast errors. It should be noted that while the weights sum to unity, some weights may be negative. Negative weights arise quite naturally, especially when the forecast errors are highly correlated (Winkler 1981), but for convenience in interpretation, we shall restrict all weights to be non-negative. The results do not change substantially if we allow for negative weights.

The forecast-error covariance matrix Σ needs to be estimated from the realized forecast errors available at the time of the formation of the combined forecast. This means that Σ needs to be estimated from a small number of observations. We use a prior distribution to stabilize the estimate of Σ (see appendix 2 for details). Before turning to the results, it should be noted that we have not conducted this weighting experiment for GDP due to the real-time problems related to this variable.

The assumption of unbiased forecast errors does not seem to hold for the DSGE and implicit-forward-rate forecasts of the interest rate (see figure 7 in appendix 2), which may have consequences for the combined forecasts. However, the combined forecast with weights inversely related to the univariate mean-squared forecast errors (MSEs), which thus include any potential bias, yields similar results in terms of its accuracy (see figure 2).

From figure 2, we also see that the combined CPI-inflation forecast performs very well at all forecast horizons. The excellent performance of the combined forecast at the first-quarter horizon is particularly noteworthy. We want to stress that we only use those forecast errors that were actually available at the time of the forecast. This means that the RMSE evaluation at, e.g., the eight-quarter horizon only uses weights up to 2003:Q4 (the sample ends in 2005:Q4).

Turning to the short-term interest rate, we see once more that combining forecasts is a good idea. The RMSE of the combined forecast is low for all forecast horizons, only slightly beaten by the implicit forward rate at the first two horizons and the BVAR forecast at the longer horizons (see the middle panel of figure 2).

3. Advantages of Structural Analysis

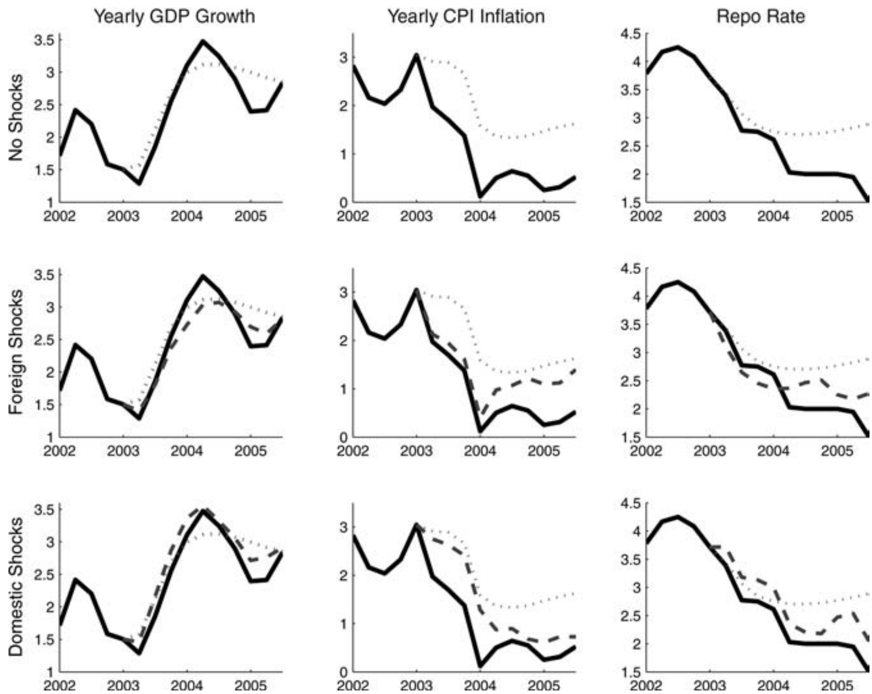
3.1 Historical Decompositions

In section 2, we analyzed the usefulness of DSGE and BVAR models for forecasting purposes. But policymakers are also very interested in understanding the factors that have brought the economy to where it is right now. This necessitates a structural model that can disentangle the underlying causes for the recent economic development. In the DSGE model, all shocks are given an economic interpretation, while the BVAR model requires additional identifying restrictions. As an example of the use of models for structural analysis, we let the two models interpret the low inflation rate in Sweden during 2003–05, by decomposing the model projections into the various shocks that have driven the development of inflation and output.

In figure 4, we report the actual outcome of GDP growth, inflation, and the repo rate together with projections from the BVAR model. The first row of figure 4 reports forecasts made in 2003:Q1 under the assumption that no shocks would hit the Swedish economy during 2003–05. It can be seen that parts of the increase in GDP growth and the decreases in inflation and the interest rate were expected, but actual inflation turned out to be a great deal lower than anticipated by the BVAR model. From the second row of figure 4—where we have added the "foreign" shocks, identified by the BVAR model ex post (dashed line), to the BVAR model's no-shock (expected) scenario (dotted line)—we can see that the sudden drop in inflation during 2003 and the lower GDP growth during the first quarters of 2003 were mainly due to foreign shocks hitting the economy.[10] The last row of figure 4 shows the effects of "domestic"

[10]The foreign shocks are identified through the assumption that foreign GDP growth, foreign inflation, and the foreign interest rate are strictly

Figure 4. Actual Outcomes (—) and Predictions for 2003:Q2–2005:Q4 from the BVAR Model without Any Shocks (···) and with Only Subsets of the Shocks Active during the Forecasting Period (− − −)



shocks, which are simply identified residually as the parts of the forecast errors that are not accounted for by foreign shocks. From the last row, we see that with only domestic shocks, the BVAR model overestimates inflation during 2003, although overall macroeconomic growth seems to be well captured. For 2004 and 2005, the picture is somewhat different, and during these years it is clear from figure 4

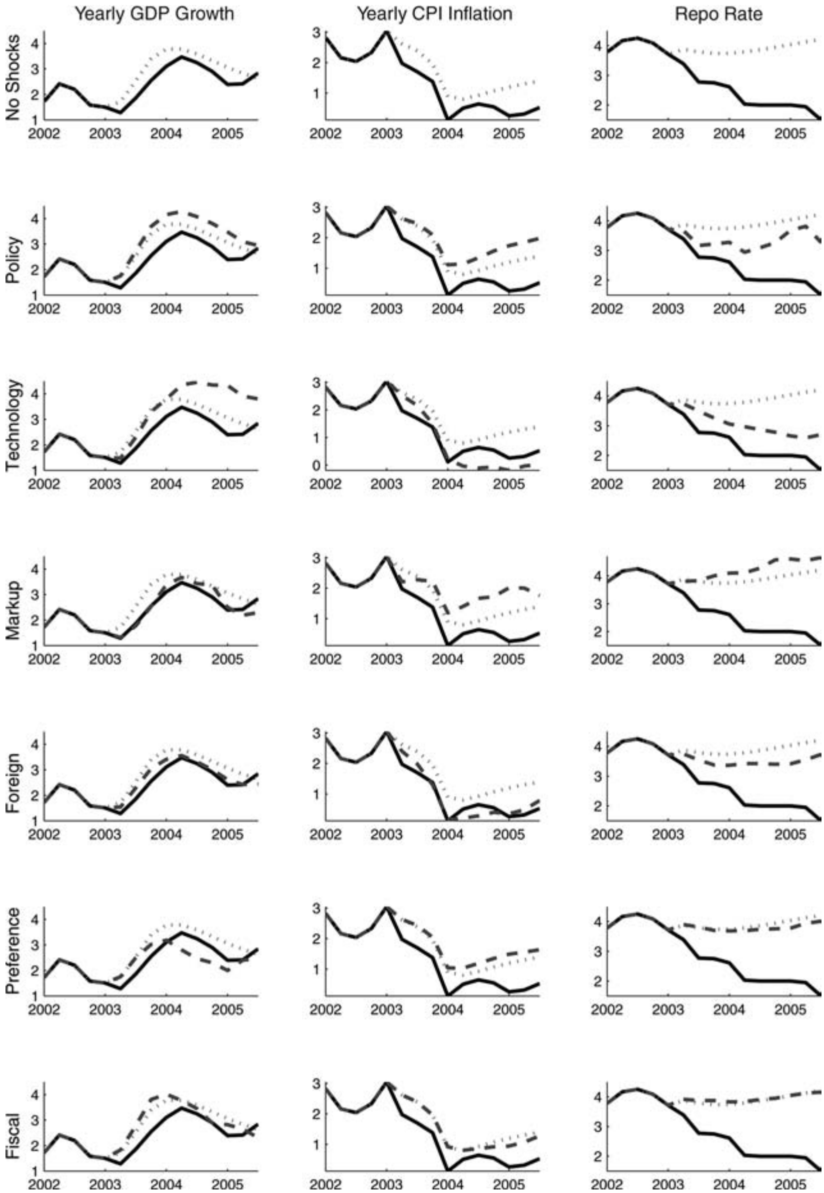
exogenous. Formally, this implies, among other things, that the foreign variables are ordered before all domestic variables in the Choleski decomposition. In addition, the forecast-error decompositions in figure 4 are based on the assumption that the real exchange rate is ordered last in the Choleski decomposition. The results are not affected much, however, if we instead assume that real-exchange-rate shocks are treated as foreign. The results are available upon request. No attempt is made to identify individual foreign shocks.

that the domestic shocks have been a more important source for the forecast errors in general and the low inflation rate in particular. In principle, it is also possible to get even more information about the shocks from the BVAR model, as long as we are willing to make additional identifying assumptions.[11]

In figure 5, we instead decompose the forecast errors from the DSGE model. The first row shows the outcome of GDP growth, inflation, and the interest rate along with the predictions from the DSGE model under the assumption that no shocks are hitting the economy (i.e., the corresponding information to the first row of figure 4, which was based on the BVAR model). Both models overestimated inflation and the interest rate, but the DSGE model also overrated GDP growth to a somewhat larger extent. The other rows in figure 5 show the "ex post forecasts" from the DSGE model when we add the model's estimates of different kinds of shocks during 2003–05 (dashed lines) to the original forecasts from 2003:Q1 (dotted lines): monetary policy shocks (second row), technology shocks (third row), markup shocks (fourth row), foreign shocks (fifth row), preference shocks (sixth row), and fiscal policy shocks (last row). It can be seen that when the estimated technology shocks and foreign shocks are individually taken into account, the "ex post forecasts" of inflation from the DSGE model are rather close to the outcome. The DSGE model thus supports the finding from the BVAR model that many of the forecast errors during 2003 were due to foreign shocks. In 2004 and 2005, when the BVAR model suggested that foreign shocks were less important, the DSGE model attributes a large part of the low inflation to both foreign shocks and domestic technology shocks. Interestingly, the model suggests that increased competition (i.e., lower markups) is not an important factor for directly understanding the low-inflation outcome. However, it is, of course, possible that the increased degree of openness (i.e., "globalization") has stimulated the favorable development in total factor productivity. It is also clear from figure 5 that fiscal policy shocks have played a very

[11] Other identifying restrictions may involve, e.g., restrictions on long-run impulses as in King et al. (1991). Jacobson et al. (2001) present some results based on such restrictions from a VAR model using similar data as this paper. Alternatively, restrictions may be imposed directly on the impulse-response functions as suggested by Canova and de Nicoló (2002) and Uhlig (2005).

Figure 5. Actual Values (—) and Predictions for 2003:Q2–2005:Q4 from the DSGE Model without Any Shocks (···) and with Only Subsets of the Shocks Active during the Forecasting Period (−−−)



limited role during this period and that monetary policy has, in fact, been expansionary. According to the DSGE model, the Riksbank reduced the repo rate more than the usual amount during this period to prevent inflation from falling too far below the target.

We find the results from these exercises very promising. Not only can the BVAR and DSGE models make forecasts that have, on average, an equal or better precision than the Riksbank's official, more subjective forecasts, but they can also, ex post, decompose the forecast errors in ways that are informative for policymakers and advisers.

3.2 The Effects of Monetary Policy Shocks

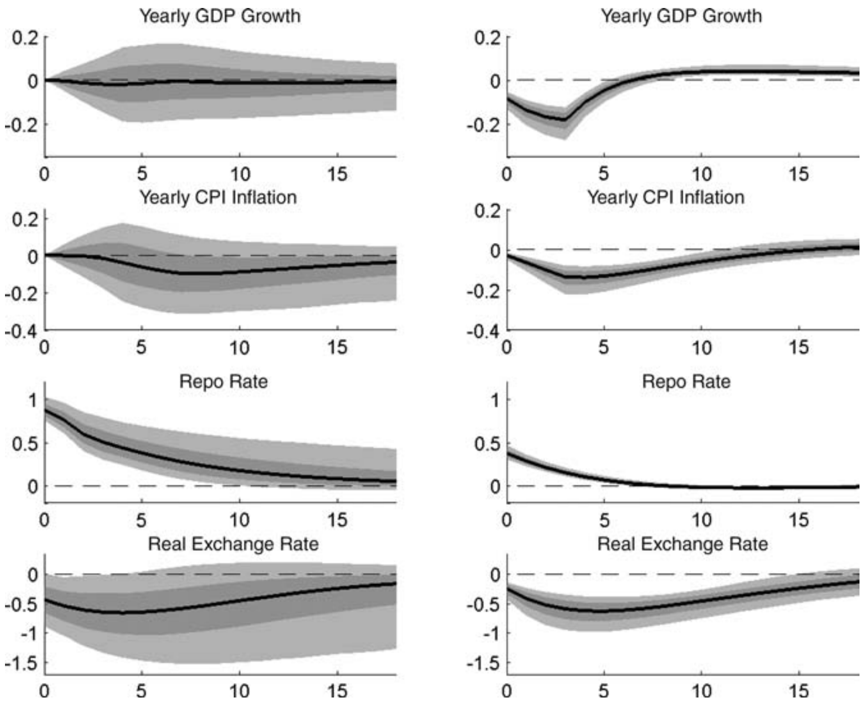
In the previous sections, we have shown that the BVAR model is a fine forecasting tool, but to some extent it can also explain what has happened in the economy, as long as we are willing to place some identifying assumptions on the shocks. However, to answer questions about how monetary policy is designed and how it influences the economy, we need to add further structure. Working with an identified model is especially important in a central bank environment where experiments such as predictions conditioned upon alternative interest rate paths are carried out. Therefore, we study impulse-response functions to see how the links between the interest rate setting and inflation outcomes differ between the BVAR and DSGE models.

Figure 6 displays the effects on output growth, CPI inflation, the repo rate, and the real exchange rate to a one-standard-deviation interest rate shock in the BVAR model (first column) and the DSGE model (last column).[12]

Qualitatively, there are some similarities between the impulse responses in the two models. For example, an increase in the interest rate appreciates the real exchange rate. However, there are also

[12]In the BVAR model, this is implemented through exogenous shocks to the interest rate in a Choleski decomposition where the interest rate is ordered after all other variables except the real exchange rate. Other nonrecursive identifying restrictions (e.g., allowing the central bank to react to changes in the real exchange rate within the period, but not to the two GDP variables) gave similar results. In the DSGE model, exogenous shocks are added to the central bank's reaction function (i.e., a monetary policy shock).

Figure 6. Posterior Median Impulse-Response Functions to a One-Standard-Deviation Interest Rate Shock for the Domestic Variables in the BVAR (First Column) and DSGE Model (Second Column) with 68 Percent and 95 Percent Probability Bands



discrepancies between the BVAR and the DSGE models. Although an increase in the interest rate causes a decline in output growth and inflation, consistent with typical prejudices, the responses in the BVAR are not significant in contrast to those in the DSGE model. Moreover, the DSGE model fulfills long-run nominal neutrality (i.e., monetary policy can only affect prices and not real quantities in the long run), whereas there is no such restriction in the BVAR. Further, the quantitative differences are very large. The DSGE model supports conventional wisdom: if the interest rate is unexpectedly increased by 0.35 percentage point (and then gradually reduced), the maximum effect on inflation is around 0.15 percentage point and is recorded after about 1–1$\frac{1}{2}$ years. The effects in the BVAR model are much smaller and typically insignificant.

Although there are reasons to expect that monetary policy shocks are more credibly identified in the DSGE model, the differences in impulse-response functions and forecasting properties between the BVAR and the DSGE models may create difficulties when using the models in the policy process, even if each individual model compares well with the subjective forecasts and rests on solid methodological grounds. Given the small impact of interest rate changes in the BVAR model, inflation projections conditional upon alternative interest rate paths would not differ to any considerable extent from the BVAR's unconditional forecasts. In contrast, the results in figure 6 suggest that the DSGE's conditional forecasts will change a great deal compared with the unconditional forecasts if the former are generated by injecting monetary policy shocks (which have much larger effects in the DSGE model). The impression of alternative policy assumptions will thus be very different in the two models.

4. Concluding Remarks

The theme of this paper is that modern macroeconomic tools like BVAR and DSGE models deserve to be used more in real-time forecasting and for policy advice at central banks. We have shown that it is possible to construct and use BVAR and DSGE models that make about as good inflation forecasts as the much more complicated judgmental procedure typically employed by central banks. In our view, central banks should use formal models—VARs and DSGEs—as benchmarks for forecasts and policy advice, and to summarize the implications of the continuous flow of new information about the state of the economy to which central bank economists are exposed.

We want to emphasize that we do not view our results as arguments against the use of judgments in monetary policy analysis. The key point here is not to dispute judgments versus formal models but, rather, to determine how to coherently combine the BVAR and DSGE models with beliefs about the current conditions. Our results suggest that it would be beneficial to incorporate judgments into the formal models, so that the forecasts reflect both judgments and historical regularities in the data.[13] We have analyzed a Bayesian weighting scheme based on the forecast errors of the different methods to combine the judgmental and model forecasts. The results are promising: the root mean-squared errors of the combined forecasts of the CPI and the interest rate are consistently low on all evaluated forecast horizons. As an alternative, short-run judgments by sector experts could be directly incorporated into the models by exploiting the methodology suggested in Waggoner and Zha (1999).

There are a number of gains in using formal models in the policy analysis. They make it possible to decompose forecast errors and provide a tool for characterizing the uncertainty involved in statements about the future development in the economy. Formal models make it possible to quantify the imprecision and uncertainties involved in the forecasting process. Formal models also serve as a learning mechanism, where lessons about the complex interdependencies in the economy can be accumulated. Our results suggest, for example, that subjective forecasts may be too myopic and not take sufficient account of important historical regularities in the data. Our present version of the DSGE model, on the other hand, may reflect problems with interest rate determination in financial markets, i.e., the empirical failure of the expectation hypothesis and the UIP condition; see the discussion in, e.g., Faust (2005).

Naturally, there are also limitations to the use of the current generation of modern macroeconomic models for policy purposes. Policymakers are often interested in details about the state of the current economy. Formal models cannot possibly cover all details within a tractable consistent framework. Neither can sector experts, but their insights into details often lead policymakers to rely on advice and forecasts from experts rather than from models. Another problem is that there are gaps between different models. Different models give quite different forecasts and imply different policy recommendations. Researchers are not typically bothered by this, as long as the models are considered to be good. Policymakers are, of course, bothered.

[13] Svensson (2005) offers a theoretical analysis of the links between judgments and monetary policy. One way of including judgments in the formal models is to approach this in a Bayesian manner. However, it is less clear how to translate the provided form of judgment into a usable prior distribution.

For some policy purposes, we therefore think that it makes sense to weight various models according to their empirical performance. We have discussed a procedure for combining forecasts that may also be used when some of the forecasts do not come from a well-specified formal model, which is typically the case in policy work. Another way of bridging the gap between formal models is to use Bayesian prior distributions that incorporate identical prior information on features that are common to the models, such as the steady state of the system (Villani 2005) or impulse-response functions (Del Negro and Schorfheide 2004).

Appendix 1. The BVAR Model

The BVAR model contains quarterly data on the following seven variables: trade-weighted measures of foreign GDP growth (y_f), CPI inflation (π_f) and the three-month interest rate (i_f), the corresponding domestic variables (y, π, and i, where i is the repo rate), and the level of the real exchange rate defined as $q = 100(s + p_f - p)$, where p_f and p are the foreign and domestic CPI levels (in logs) and s is the (log of the) trade-weighted nominal exchange rate.

The BVAR model used in this paper is of the form

$$\Pi(L)(x_t - \Psi d_t) = A\varepsilon_t, \tag{1}$$

where $x = (y_f, \pi_f, i_f, y, \pi, r, q)'$ is an n-dimensional vector of time series, $\Pi(L) = I_n - \Pi_1 L - \ldots - \Pi_k L^k$, and L is the usual back-shift operator with the property $Lx_t = x_{t-1}$. The structural disturbances $\varepsilon_t \sim N_n(0, I_n)$, $t = 1, \ldots, T$, are assumed to be independent across time. We impose restrictions on $\Pi(L)$ such that the foreign economy is exogenous. A is the lower-triangular (Choleski) contemporaneous-impact matrix, such that the covariance matrix Σ of the reduced-form disturbances decomposes as $\Sigma = AA'$. We also have experimented with nonrecursive identifying restrictions, in which case the equations are normalized with the Waggoner-Zha rule (Waggoner and Zha 2003b), and the Gibbs sampling algorithm in Waggoner and Zha (2003a) is used to sample from the posterior distribution. The deterministic component is $d_t = (1, d_{MP,t})'$, where

$$d_{MP,t} = \begin{cases} 1 & \text{if } t < 1993\text{:Q1} \\ 0 & \text{if } t \geq 1993\text{:Q1} \end{cases}$$

is a shift dummy to model the abandonment of the fixed exchange rate and the introduction of an explicit inflation target in 1993:Q1. Since the data are modeled on a quarterly frequency, we use $k = 4$ lags in the analysis. Larger lag lengths gave essentially the same results, with a slight increase in parameter uncertainty.

The somewhat nonstandard parameterization of the VAR model in (1) is nonlinear in its parameters but has the advantage that the unconditional mean, or steady state, of the process is directly specified by Ψ as $E_0(x_t) = \Psi d_t$. This allows us to incorporate prior beliefs directly on the steady state of the system, e.g., the information that steady-state inflation is likely to be close to the Riksbank's inflation target. To formulate a prior on Ψ, note that the specification of d_t implies the following parameterization of the steady state:

$$E_0(x_t) = \begin{cases} \psi_1 + \psi_2 & \text{if } t < 1993\text{:Q1} \\ \psi_1 & \text{if } t \geq 1993\text{:Q1}, \end{cases}$$

where ψ_i is the i-th column of Ψ. The elements in Ψ are assumed to be independent and normally distributed a priori. The 95 percent prior probability intervals are given in table 1.

The prior proposed by Litterman (1986) will be used on the dynamic coefficients in Π, with the default values on the hyperparameters in the priors suggested by Doan (1992): overall tightness is set to 0.2, cross-equation tightness to 0.5, and a harmonic lag decay with a hyperparameter equal to 1. See Litterman (1986) and Doan (1992) for details. Litterman's prior was designed for data in levels and has the effect of shrinking the process toward the univariate random-walk model. Therefore, we set the prior mean on the first own lag to 0 for all variables in growth rates. The two interest rates and the real exchange rate are assigned a prior that centers on the $AR(1)$ process with a dynamic coefficient equal to 0.9. The usual

Table 1. Ninety-Five Percent Prior Probability Intervals of Ψ

	y_f	π_f	r_f	y	π	r	q
ψ_1	(2, 3)	(1.5, 2.5)	(4.5, 5.5)	(2, 2.5)	(1.7, 2.3)	(4, 4.5)	(−1, 1)
ψ_2	(−1, 1)	(1.5, 2.5)	(1.5, 2.5)	(−1, 1)	(4.3, 5.7)	(3, 5.5)	(−9, 9)

random-walk prior is not used here, as it is inconsistent with having a prior on the steady state. Finally, the usual non-informative prior $|\Sigma|^{-(n+1)/2}$ is used for Σ.

The posterior distribution of the model's parameters and the forecast distribution of the seven endogenous variables were computed numerically by sampling from the posterior distribution with the Gibbs sampling algorithm in Villani (2005).

Appendix 2. Combining Judgmental and Model Forecasts

Suppose that we have available forecasts, at a given forecast horizon, from k different forecasting methods over T different time periods. Let $\hat{x}_{jt}$ denote the j-th method's forecast of a variable x_t, and $e_{jt} = \hat{x}_{jt} - x_t$ the corresponding forecast error, where $j = 1, \ldots, k$ and $t = 1, \ldots, T$. The question here is how to merge these k forecasts into a single combined forecast. Following Winkler (1981), we shall assume that the vector of forecast errors from the k methods, $e_t = (e_{1t}, \ldots, e_{kt})'$, can be modeled as independent draws from a multivariate normal distribution with zero mean and covariance matrix Σ. This implies that $\hat{x}_t \sim N_k(x_t u, \Sigma)$, where $\hat{x}_t = (\hat{x}_{1t}, \ldots, \hat{x}_{kt})'$ is the vector of forecasts of x_t from the k forecasting methods, and $u = (1, \ldots, 1)'$. We use an uninformative (uniform) prior on x_t and an inverted Wishart density for Σ a priori: $\Sigma \sim IW(\Sigma_0, v)$, where $\Sigma_0 = E(\Sigma)$ and $v \geq k$ is the degrees-of-freedom parameter. The prior on Σ is important, as historical forecast errors are limited and an estimate of Σ is typically unreliable. This is particularly important when the correlations in Σ are large, which is often the case with forecast errors from competing methods. As v increases, the prior becomes increasingly concentrated around Σ_0. The specification of Σ_0 and v is discussed below.

The posterior mean of the true value x_t, which is the natural combined forecast for a Bayesian, can now be shown to be a linear combination of the individual forecasts (Winkler 1981):

$$E(x_t|\hat{x}_t) = \sum_{j=1}^{k} w_{jt}\hat{x}_{jt}, \tag{2}$$

where the weights of the forecasting methods are given by

$$w_t' = (w_{1t}, \ldots, w_{kt}) = \frac{u'\tilde{\Sigma}_t^{-1}}{u'\tilde{\Sigma}_t^{-1}u}, \tag{3}$$

and the posterior estimate of Σ is

$$\tilde{\Sigma}_t = E(\Sigma|e_1, \ldots, e_t) = \frac{v}{t+v}\Sigma_0 + \frac{t}{t+v}\hat{\Sigma}_t,$$

where $\hat{\Sigma}_t$ is the usual unbiased estimator of a covariance matrix.

The weights w_t sum to unity at all dates t, but they need not be positive. Negative weights may result quite naturally, as explained in Winkler (1981), especially when the forecasts are positively correlated across methods.

Strictly speaking, the weighting scheme in (3) is only known to be the Bayesian solution under the assumption that forecast errors are independent and unbiased. The independence assumption is likely to be violated for forecast errors beyond the first horizon. For simplicity, we will continue to assume independent forecast errors at all forecast horizons. An alternative approach would be to stick to the weighting scheme in (3) but with a more sophisticated $\hat{\Sigma}_t$ estimate that accounts for autocorrelation, e.g., the Newey-West estimator. This procedure is unlikely to be a Bayesian solution, however, and also suffers from the drawback that the Newey-West estimator is likely to be unstable when the history of available forecast errors is short. The second assumption behind (3) is that forecasts are unbiased. This does not seem to be supported for the implicit forward rate or the DSGE's interest rate forecast at longer horizons, and can potentially have a large effect on the combined forecast. Therefore, we also look at an ad hoc method for combining forecasts with weights inversely proportional to the mean-squared errors from past forecasts. Note that this method ignores the fact that forecast errors of different methods are typically correlated.

We need to determine Σ_0 and v in the inverted Wishart prior for Σ. We will use the following parameterization of the prior mean of Σ:

$$\Sigma_0 = \sigma_0^2 \begin{pmatrix} 1 & \rho & \rho \\ \rho & 1 & \rho \\ \rho & \rho & 1 \end{pmatrix},$$

which leaves σ_0, ρ, and v to be specified. It seems fair to expect all forecasting methods to produce fairly correlated forecasts, so that ρ is comparatively large and also increases with the forecast horizon (most methods will produce long-run forecasts that are quite close to their steady-state value, and the steady states in the different methods should not be too different). Moreover, σ_0 should also increase with the forecast horizon. We will assume that both ρ and σ_0 increase linearly with the forecast horizon (σ_0 ranges from 0.5 to 1, and ρ equals 0.5 at the first forecast horizon and 0.92 at the eighth horizon). Finally, we need to pin down the overall precision in the prior—the degrees-of-freedom parameter, v. We set $v = 50$, which gives us a 95 percent prior probability interval for ρ at the first horizon equal to (0.4, 0.75). The results are robust to nondrastic variations in the prior.

Turning to the results, we show the bias for the different models' CPI inflation and interest rate forecasts in figure 7. The interest rate forecasts from the DSGE model and the implicit forward rate especially seem to be biased. Therefore, we also consider a weighting scheme that is inversely related to the univariate mean-squared

Figure 7. Forecast Bias

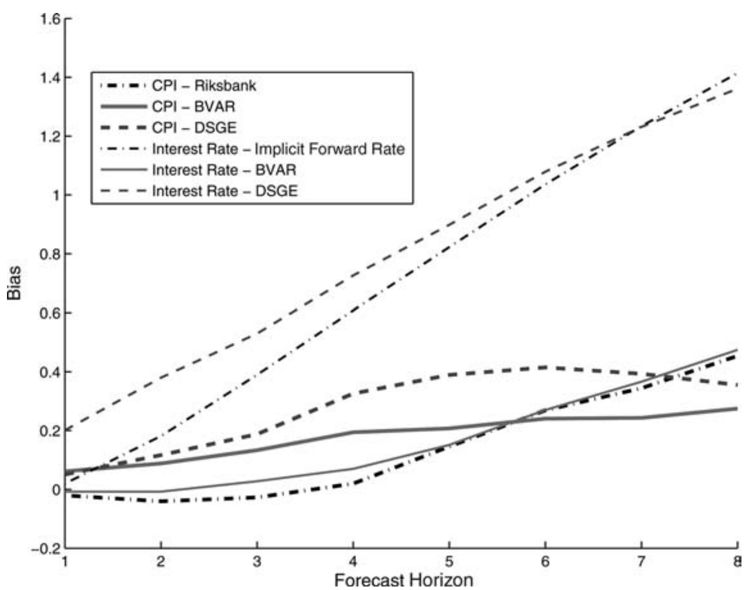
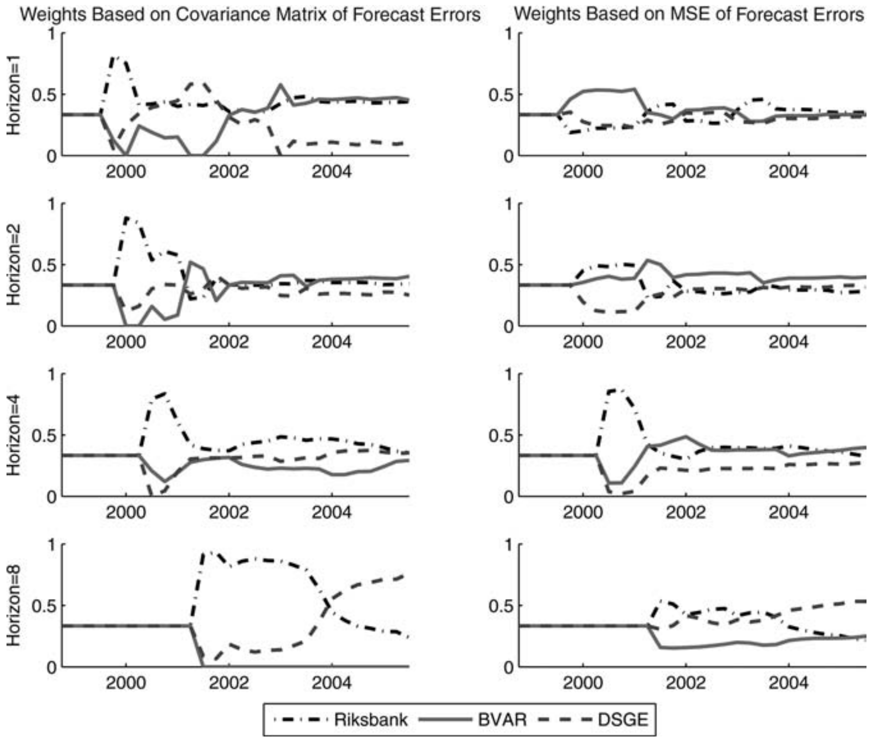


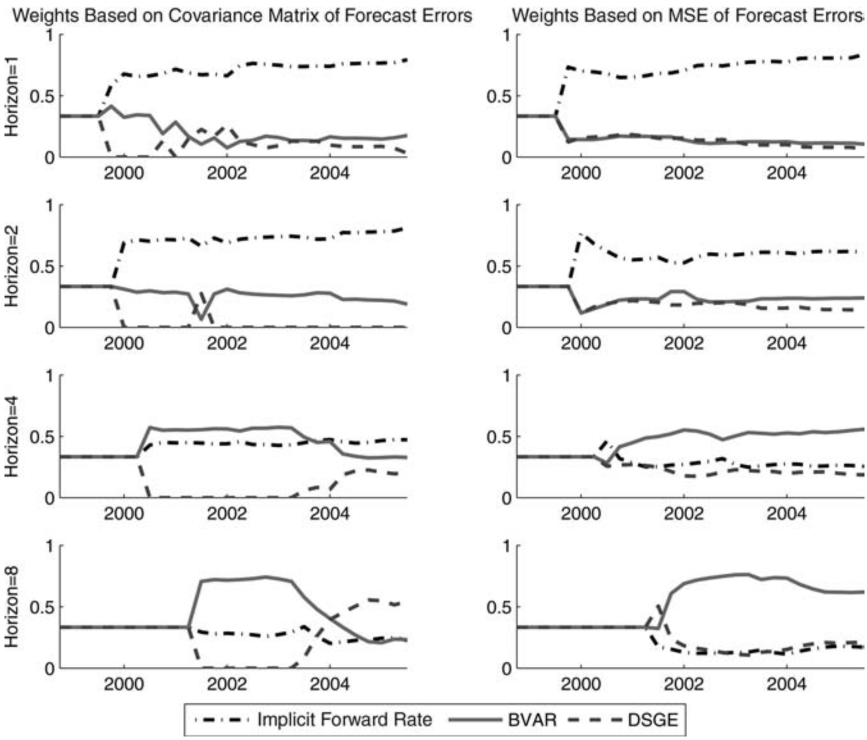
Figure 8. Sequential Weighting Schemes for Yearly CPI Inflation



forecast errors, which thus includes any potential bias. In figures 8 and 9, we compare the two weighting schemes. However, as seen from figure 2 in the main text, the accuracy of the combined forecast does not seem to be affected much by which scheme is used.

Figure 8 shows the sequential weighting schemes for CPI inflation at different horizons. Note that the weights are equal on all three forecasts at the beginning of the evaluation period, where there are not enough realized forecast errors to estimate Σ. The weights for CPI inflation in the two different weighting schemes are similar on the first- and second-quarter horizons, at least in the latter part of the evaluation period. At longer horizons, there are fewer forecast errors for constructing the weights and larger biases in the forecasts. This causes the two weighting schemes to differ much more than at shorter horizons.

Figure 9. Sequential Weighting Schemes for the Interest Rate



For the interest rate, the sequential forecast weights are once more stable at the two shortest horizons compared with the longer ones; see figure 9. From figure 9, it can also be seen that the superiority of the implicit forward rate at the first- and second-quarter horizons is immediately picked up by both weighting schemes.

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A New Core Inflation Indicator for New Zealand*

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This paper introduces a new indicator of core inflation for New Zealand, estimated using a dynamic factor model and disaggregate consumer price data. Using disaggregate consumer price data, we can directly compare the predictive performance of our core indicator with a wide range of other "core inflation" measures estimated from disaggregate consumer prices, such as the weighted median and the trimmed mean. The medium-term inflation target of the Reserve Bank of New Zealand is used as a guide to define our target measure of core inflation—a centered two-year moving average of past and future inflation outcomes. We find that our indicator produces relatively good estimates of this characterization of core inflation when compared with estimates derived from a range of other models.

JEL Codes: C32, E31, E32, E52.

1. Introduction

Inflation is a key variable in the formulation of monetary policy. However, inflation data are often subject to transitory shocks, which do not require a policy response. The purpose of core inflation measures is thus to remove the influence of transitory shocks, revealing the unobserved, policy-relevant trend in inflation.

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The 2002 Policy Targets Agreement (PTA), signed by the Minister of Finance and the governor of the Reserve Bank of New Zealand (RBNZ), requires the Reserve Bank to keep annual inflation in the consumers price index (CPI) between 1 and 3 percent, on average, over the medium term. This suggests that the policy-relevant rate of inflation for New Zealand should remove short-term fluctuations from inflation by averaging over the medium term. But what is the "medium term"?

It is commonly thought that the lag between a change in monetary policy and its impact on inflation is one to three years, suggesting that the medium term is a period long enough to smooth out fluctuations in annual inflation lasting around two years in duration.[1] Core inflation could thus be defined as historical realizations of the inflation target, where the target looks something like an average of annual inflation over the past two years. The problem with this characterization of core inflation, however, is that it lags annual CPI inflation by almost a year.

Ideally, we would like our core inflation indicator to look like a moving average (MA) of annual inflation, but without the phase shift induced by taking an average of historical data. In other words, we want an indicator that looks like a *centered* moving average of annual inflation (or, equivalently, a forecast of a backward-looking average), but one that can be computed in real time. One way or another, this is precisely what all core inflation measures aim to do; they aim to produce a real-time estimate of the underlying, policy-relevant rate of inflation, something that is unobservable in real time.

We estimate core inflation using disaggregate consumer prices and a cross-sectional smoothing technique that was first used by Altissimo et al. (2001) to estimate a coincident indicator for the euro-area business cycle. Essentially, we aim to exploit the leading and lagging relationships among disaggregate prices to proxy for future CPI inflation, thus smoothing inflation and avoiding the phase shift induced by taking an average of historical data on aggregate prices.

Many central banks monitor a variety of measures of core inflation, typically constructed from disaggregate CPI data. The CPI

[1] Christiano, Eichenbaum, and Evans (2005), e.g., find that inflation responds in a hump-shaped fashion to a monetary shock, peaking at around about two years.

excluding food and energy, for example, is a very popular measure of core inflation in the United States. There is a multitude of other so-called exclusion measures used around the world, all with a common goal in mind—to eliminate the idiosyncratic noise from inflation by removing the most volatile components from the CPI. In addition to the exclusion measures, there is also a range of statistical measures of core inflation monitored by central banks. These measures aim to remove those components from inflation that are most volatile in a particular month, quarter, or year. Two popular statistical measures are the weighted median and the trimmed mean of inflation, proposed by Bryan and Cecchetti (1994). However, many of these measures of core inflation—both the exclusion and the statistical methods—tend to be too volatile and often fail to provide a reliable signal for underlying inflation (Cogley 2002).

Recently, there have been some advances in econometric theory that allow the decomposition of very large panels of data into a small number of common factors (Forni et al. 2000; Stock and Watson 2002; Forni et al. 2005). These methods are well suited to the problem of estimating core inflation, where the inflation signal of interest is both unobservable in real time and common to a large number of macroeconomic series.

Cristadoro et al. (2005) apply the dynamic factor model to extract the long-run component of inflation from almost 450 nominal, real, and financial indicators of inflation, producing a measure of core inflation for the euro area. Their core measure has the benefit of being smooth, without having the phase shift induced by taking an annual percentage change of inflation. Moreover, the Cristadoro et al. (2005) indicator is very competitive at forecasting annual inflation at horizons up to two years ahead. Similarly, Amstad and Fischer (2004) and Amstad and Potter (2007) use the dynamic factor model to compute core inflation measures for Switzerland and the United States, respectively. Unlike Cristadoro et al. (2005), Amstad and Fischer (2004) and Amstad and Potter (2007) allow for real-time estimates of core inflation as each new piece of data arrives (what the authors call "sequential information flow").

One of the key objectives of this paper is to compare dynamic factor model estimates of core inflation with a variety of methods of estimating core inflation from disaggregate consumer price data—including standard core inflation measures, estimates from pooling regressions, and estimates from bivariate forecasting models. Essentially, we want to ask, Given data on disaggregate consumer price movements, what is the best way to estimate core inflation? Thus, unlike Amstad and Fischer (2004), Cristadoro et al. (2005), and Amstad and Potter (2007) (who use broader data sets to estimate core inflation), we limit ourselves to using disaggregate consumer prices.[2]

We assess the performance of our core inflation indicator in real time by comparing how well it predicts a particular characterization of the RBNZ's inflation target—a centered two-year moving average of annual inflation. We compare the indicator with a variety of other estimates of core inflation based on disaggregate price data and find that the indicator performs well. Furthermore, when compared with a range of forecasting models used at the RBNZ, many of which use a wider range of information, the core inflation measure also compares favorably and is only bettered by models that incorporate judgment and utilize more up-to-date information.

2. Methodology

We have T time-series observations for N different inflation series from the CPI denoted π_{jt}, where $j = 1, \ldots, N$; $t = 1, \ldots, T$; and π_{jt} is the (log) seasonal change of the j-th price index. Further, let us add to this panel the (log) seasonal change in headline CPI, π_t. Headline inflation can then be represented as the sum of two unobserved components, a signal π_t^* and an error e_t:

$$\pi_t = \pi_t^* + e_t. \tag{1}$$

The objective is to estimate the signal π_t^* using all information in the panel of CPI price changes. We assume that each variable can be represented as two stationary, orthogonal, unobservable components—a common component χ_{jt} and an idiosyncratic component ε_{jt}:

$$\pi_{jt} = \chi_{jt} + \varepsilon_{jt}, \tag{2}$$

[2]In the case of New Zealand, the CPI data are not subject to revision and are more timely than most other key macroeconomic releases; the CPI is published more than one month before the producer price index and more than two months before the national accounts.

where the common component is driven by a small number of common factors (shocks).

We decompose the common component into a long-run component χ_{jt}^{L} and a short-run component χ_{jt}^{S} by removing high-frequency, short-run fluctuations up to a given critical period h (Cristadoro et al. 2005):

$$\pi_{jt} = \chi_{jt}^{L} + \chi_{jt}^{S} + \varepsilon_{jt}. \tag{3}$$

Specifically, the intertemporally smoothed (long-run) common component can be attained by summing waves of different periodicity between $[-\pi/h, \pi/h]$ using a spectral decomposition. The long-run common component is what we are after in estimating our measure of core inflation. This measure removes the idiosyncratic noise specific to each of the components of the CPI, as well as smoothing out the short-term fluctuations not requiring a monetary policy response.

Isolation of the unobserved common component can be achieved by assuming that the common components are driven by shocks that are pervasive in the cross-section of price movements, while the shocks driving the idiosyncratic terms are local and affect only a limited number of prices. This is called an approximate dynamic factor structure. The dynamic factor model is

$$\pi_{jt} = b_j(L) f_t + \varepsilon_{jt}, \tag{4}$$

where $f_t = (f_{1t}, \ldots, f_{qt})'$ is a vector of q dynamic factors and $b_j(L)$ is of order s, for every dynamic factor $1, \ldots, q$: the process for f_t is assumed to be stationary.[3] This model is said to have q common dynamic factors.

If we let $F_t = (f_t', f_{t-1}', \ldots, f_{t-s}')'$, the dynamic factors have a static representation:

$$\pi_{jt} = \lambda_j F_t + \varepsilon_{jt}, \tag{5}$$

where $b_j(L) f_t = \lambda_j F_t$. Thus, a model with q dynamic factors has $r = q(s+1)$ static factors.

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[3]Typically the data-generating process for the common factors is approximated by a finite-order vector autoregression (VAR). See, e.g., Forni et al. (2007).

3. Estimation

The dynamic factor model is estimated in the frequency domain using an eigenvalue decomposition of the spectrum smoothed over a range of frequencies. Estimation requires the specification of three parameters: two of the three parameters determining the number of static factors r, q, and s, and the size of the Bartlett lag window M.[4] Estimation of the static factor representation of the dynamic factor model (5), on the other hand, requires an eigenvector-eigenvalue decomposition of the variance-covariance matrix and requires r to be specified.[5] We use the dynamic factor representation to estimate our indicator of core inflation. More details on the estimation of the dynamic and static factor models are provided in appendix 1.

In order to get good estimates of the common factors, Cristadoro et al. (2005) recommend that the panel of data used to estimate the dynamic factor model should have series that lead and lag annual inflation. Their argument is summarized as follows.

Consider the case where there is only one common shock to inflation, f_t, which is loaded with different lags in a cross-section containing three variables:

$$\chi_{1t} = f_{t-1}, \quad \chi_{2t} = f_t, \quad \text{and} \quad \chi_{3t} = f_{t-2}. \tag{6}$$

In this case, it is clear that variable 2 leads variable 1, which in turn leads variable 3. If CPI inflation is χ_{1t}, then $\chi_{1t+1} = f_t$ and $\chi_{1t-1} = f_{t-2}$, so that the cross-sectional information hidden in the contemporaneous $\chi_{jt}s$ is exactly the time-series information contained in CPI inflation and its first lead and lag. CPI inflation can thus be smoothed in period t by using the leading variable as a proxy for future CPI inflation, which is unavailable at time t.

The methodology then consists of projecting the long-run common component of headline CPI inflation, χ_t^L, onto the (intertemporally smoothed) present and past common factors:

$$\hat{\chi}_t^L = \text{Proj}\big[\chi_t^L \big| f_{mt-k}, m = 1, \ldots, q; k = 1, \ldots, s\big]. \tag{7}$$

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[4]For all the dynamic factor models estimated in this paper, we set $M = \sqrt{T}$, as in Forni et al. (2005).

[5]A VAR in F_t can then be used to estimate the dynamics of the model.

Thus, as seen in the example above, the method will produce good results if χ_t loads mainly with lags central to the interval $0, \ldots, s$. In other words, to get good estimates of the common factors, the data set must include variables that lead and lag χ_t. As noted by Cristadoro et al. (2005), leading variables are of particular importance, since lagging variables could be replaced by lags of existing variables, whereas the leading variables are irreplaceable.

As noted in the introduction, we characterize the target for our core inflation indicator as a centered two-year moving average of annual CPI inflation. Letting $MA(\pi_t, h)$ be a centered moving average with a window of $2h + 1$, then

$$MA(\pi_t, h) = \frac{1}{2h+1}(\pi_{t-h} + \pi_{t-h+1} + \ldots + \pi_t + \ldots + \pi_{t+h-1} + \pi_{t+h}). \tag{8}$$

Our objective is to construct an indicator of core inflation that matches the properties of this filter and does not require future information (information from period $t + 1$ to $t + h$) to compute. Estimation in the frequency domain allows us to remove short-term fluctuations from the common components to reveal a long-run common component of annual CPI inflation $\hat{\chi}_t^L$. Moreover, our target filter $MA(\pi_t, 4)$ sets the long-run frequencies that we can use to intertemporally smooth the common component $\hat{\chi}_t^L$. Specifically, we compute the intertemporally smoothed (long-run) common component of inflation by summing waves of different periodicity in the band $[0, 2\pi/9]$, removing all short-run cyclical fluctuations up to nine quarters in duration.[6] Henceforth, our characterization of the inflation target is denoted $\pi_t^{target} = MA(\pi_t, 4)$ and the dynamic factor model's estimate of core inflation is denoted $\pi_t^{core} = \hat{\chi}_t^L$.

4. Data and Dynamic Structure of the Data

We begin with quarterly data for all 264 subsections of the CPI for a period ranging from 1991:Q1 to 2006:Q2. To these series we also add headline CPI, tradable CPI, and nontradable CPI.[7]

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[6] See Giannone and Matheson (2006) for further discussion.

[7] Statistics New Zealand publishes a split of the CPI regimen into tradable and nontradable price indexes. We include these indexes in our panel to enable

The data are filtered in five steps. First, we remove all series from panel that do not span the entire sample. Second, we take the natural logarithm and seasonally difference all series to achieve stationarity. Third, we remove outliers from each series by replacing observations more than six times the interquartile range with the median of the series. Fourth, we remove the series whose prices change less than once a year on average.[8] The filtered series are then standardized to have zero mean and a unit variance. After filtering, the panel contains headline inflation, tradable inflation, nontradable inflation, and inflation data for 228 subsections of the CPI.[9] Core inflation is obtained by reattributing the mean and the variance of each series; i.e., the estimate of π_t^{core} is rescaled by multiplying it by the standard deviation of π_t and by adding the mean of π_t.

The left side of figure 1 displays a histogram of the leads/lags at which the highest absolute correlation with inflation occurs in our panel. The figure shows that the structure of the data is well balanced and includes a similar number of leading and lagging variables. Because of the importance of the leading variables in estimation, we also examine the predictive power of the series in our panel by way of a Granger causality test—testing the null hypothesis that series j does not Granger-cause headline CPI inflation. The right side of figure 1 displays a histogram of the p-values from this test. Around 30 percent of the series in the panel Granger-cause headline inflation at the 10 percent level of significance. With a similar number

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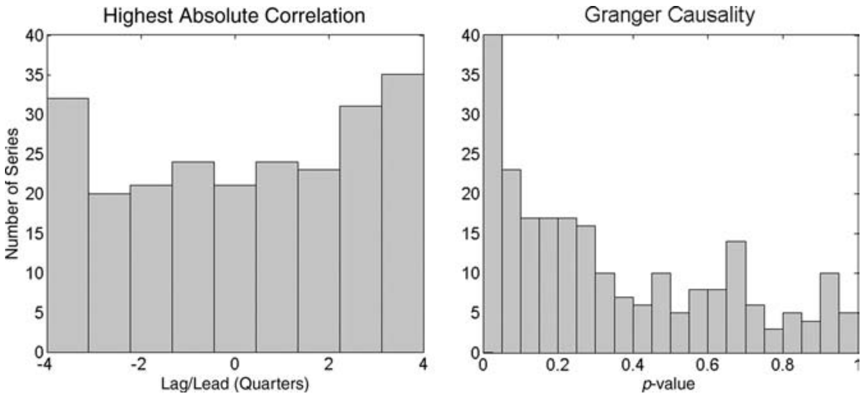
the RBNZ to more readily compute core measures of tradable and nontradable inflation. The inclusion of these aggregate variables should not pose a problem for estimating core inflation, because an approximate factor structure allows for local correlation among the idiosyncratic components.

[8] Correcting for outliers and removing series that change less than once a year on average are not crucial steps. The core indicator estimated with the raw, unadjusted data is very similar to the indicator estimated with the adjusted data: the root mean-squared forecast error of the core indicator estimated without adjustments is only marginally (0.0052 percent) higher than that presented in table 4 (in section 6). Nevertheless, we choose to make these corrections to insure against any coding or typographical errors (made by the RBNZ or Statistics New Zealand) that could adversely affect quarterly updates of core inflation in the future. The core inflation indicator has been published in the RBNZ's quarterly *Monetary Policy Statement* since September 2006.

[9] During the filtering process, a total of 9.21 percent of the CPI regimen is removed from the original panel and thirty-five observations are classified as outliers. See Giannone and Matheson (2006).

Figure 1. The Correlation Structure of the Panel



of leading and lagging variables and with a sizable proportion of the series having some predictive ability with respect to annual CPI inflation, this panel seems well suited to estimating a dynamic factor model. A more detailed description of our panel, including statistics relating to the dynamic structure of the data discussed in the next section, can be found in Giannone and Matheson (2006).

4.1 The Dynamic Structure of the Data

Determining the unobserved factors driving prices is difficult. As with many statistical problems, the choice of the number of factors requires a trade-off between parsimony and fit. The more factors that are added to the model, the more variation in the data set that will be explained by the factors. Fewer factors, on the other hand, will produce a smoother indicator of core inflation, but at a cost of poorer fit to the data. Moreover, the selection of the number of static factors r is further complicated since there is more than one configuration of the parameters q and s that has similar fit to the data.

Bai and Ng (2002) show that the number of static factors can be consistently estimated with an information criterion. Unfortunately, the Bai and Ng criterion does not appear to be suitable for our empirical application and chooses T static factors. Effectively, this means that, statistically, our panel is driven by very many factors—so many, in fact, that if one were to project inflation on the estimated common components, one would get precisely the original data back, rotated on the factors.

In an approximate factor structure with q common shocks and r common factors, there are q dominant eigenvalues from the spectral-density matrix (dynamic rank) and r dominant eigenvalues from the covariance matrix (static rank). Thus, some insight into a number of the common shocks can be found by examining the behavior of the dynamic eigenvalues, which represent the variance of the dynamic factors. Figure 2 plots the first ten dynamic eigenvalues from the spectral-density matrix of our data over frequencies $[0, \pi]$. The first four eigenvalues are considerably larger than the others, particularly at the long-run frequencies with which we are concerned. It thus seems that the long-run common comovement in these data can be adequately captured by the first four dynamic factors.

Another approach to determine the number of factors is to describe the comovements of the panel using the percentage of the

Figure 2. The First Ten Dynamic Eigenvalues from the Spectral-Density Matrix

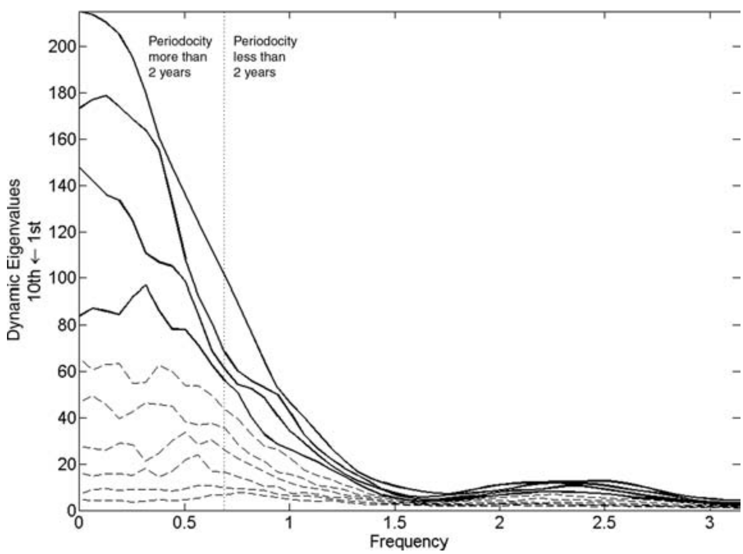


Table 1. Percentage of Total Variance Explained by the First Twelve Dynamic and Static Factors

	1	**2**	**3**	**4**	**5**	**6**	**8**	**10**	**12**
q	0.23	0.42	0.57	0.69	0.78	0.84	0.92	0.95	0.97
r	0.13	0.26	0.35	0.42	0.48	0.53	0.63	0.70	0.76

variance of the panel accounted for by the common factors, as suggested by Forni et al. (2000). Table 1 reports the percentage of the total variance in our panel explained by the first twelve dynamic factors q (estimated using dynamic principal components) and the same number of static factors r (estimated using principal components).[10] The table shows that a small number of factors explain a large amount of the variation in our panel. Moreover, there is a discrepancy between the variance explained by the dynamic and static factors, suggesting, as with the results in section 4, that there are some rich dynamics at play in our panel. Because the rank of the covariance matrix of the panel is always $r = q(s+1)$ and the rank of the spectral-density matrix is always q, the difference between the variance explained by r and q reflects the lagged factors s. Selecting the number of dynamic factors q so that the marginal contribution from adding one more factor is less than 10 percent, as suggested by Forni et al. (2000), produces $q = 4$. Selecting the number of static factors r to explain the same amount of variation as $q = 4$ produces $r \approx 10$, implying that s is somewhere between 1 and 2.

For our indicator of core inflation, we choose to set $q = 4$, $s = 2$, and $r = 12$.[11] As a robustness check, we estimate the indicator recursively with different configurations of q and s over a period from 2000:Q1 to 2005:Q2 and find that our chosen parameterization performs comparatively well. Indeed, with the exception of when $s = 0$, the models with $q = 4$ outperform all other models for a given s, justifying our assumption of four dynamic factors. The

[10] See appendix 1 for a description of principal-components estimators.

[11] Equation (5) implies that $r = q(s + 1)$. However, it is worth noting that this holds only in the case where the order of the MA in the common shocks is finite and that, in general, $r \geq q(s + 1)$; see Forni et al. (2007). In choosing our parameter configuration, we implicitly assume that the order of the MA is finite.

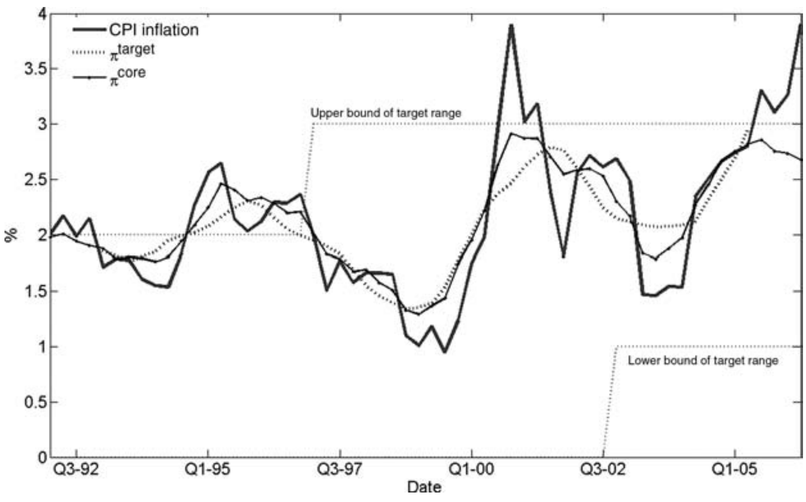
performance of the indicator with $q = 4$ and $s = 1$ is the same as our chosen parameterization.[12]

Giannone and Matheson (2006) show that the first four common factors explain more than 75 percent of the variability of headline CPI inflation. Moreover, at the periodicities with which we are concerned (longer than two years), the degree of commonality is even higher, with the common factors explaining over 80 percent of the variation.

5. The Core Inflation Indicator

The core inflation indicator is displayed in figure 3 alongside annual CPI inflation and our target variable, a centered two-year moving average of annual CPI inflation. We find that the core inflation indicator smooths much of the noise from annual CPI inflation, and it closely tracks the target variable.

Figure 3. CPI Inflation, the Inflation Target, and the Core Indicator



[12]See section 6 for more details on the real-time prediction experiment and appendix 4 for the results for different configurations of q and s.

So how well does our core inflation indicator do in predicting the target measure of inflation? Table 2 compares the core inflation indicator with weighted-median inflation; trimmed-mean inflation (with a 10 percent trim); the CPI excluding food, administration changes, and gasoline; and the exponentially smoothed measure of core inflation proposed by Cogley (2002). All of these measures of core inflation are described in greater detail in appendix 2 of the paper. For the moment, we only discuss our core inflation indicator estimated up to the end of the sample. The real-time indicator in the final row of the table, Core (Real Time), is discussed in the next section.

All of the core measures average around 2 percent, similar to headline CPI. However, there are some large differences in the variability of the measures. In particular, the measures of the weighted median, trimmed mean, and CPI excluding food, administration charges, and gasoline are much more volatile than our core indicator and the exponentially smoothed indicator. In fact, these core measures seem to do a bad job at smoothing inflation, having standard deviations higher than headline CPI inflation itself. Our core indicator has a standard deviation that closely matches the target variable, whereas the exponentially smoothed measure seems to smooth inflation too much and has a much lower standard deviation than the target.[13]

Looking at absolute correlations with the target measure, π_t^{target}, at leads and lags of up to one year, we find that our core inflation indicator correlates highly and is in phase with the target variable. The exponentially smoothed measure also correlates highly with the target, although with a lag of two quarters. Likewise, the weighted median, the trimmed mean, and the CPI excluding food, administration charges, and gasoline tend to lag the target by a couple of quarters.

It is interesting to look at the concordance of each core inflation measure with our target variable, π_t^{target}. Concordance measures the percentage of the sample where changes to the indicator and changes to the target variable have the same sign, i.e., the percentage of time that changes to the core indicator accurately reflect changes in the

[13]The standard deviation of π_t^{target} is 0.46.

Table 2. Descriptive Statistics for Various Measures of Core Inflation

	Mean	Std.	Max.	Lag	Conc.	$\pi_{t+4} - \pi_t = \alpha + \beta(\pi_t^{core} - \pi_t) + \epsilon_{t+4}$		
						α	β	R^2
CPI	2.07	0.72	0.77	0	0.70	—	—	—
Core	2.14	0.51	0.95	0	0.79	−0.06	2.58*	0.71
Weighted Median	2.22	0.79	0.82	3	0.55	0.13	0.03	0.00
Trimmed Mean	2.08	0.76	0.83	1	0.58	0.13	−0.27	0.01
CPI excl. Food, Admin., and Gasoline	1.89	0.88	0.83	2	0.64	0.05	−0.45	0.04
Exponentially Smoothed	2.04	0.29	0.94	2	0.70	0.11	1.17*	0.38
Core (Real Time)	2.03	0.32	0.92	1	0.82	0.13	1.31*	0.44

Notes: Mean of the series (Mean); Standard deviation (Std.); Maximum absolute correlation with π_t^{target} (Max.); Lag at which the maximum absolute correlation occurs (Lag); Concordance with π_t^{target} (Conc.); coefficient on the deviation from each core measure to predicting $\pi_t + 4 - \pi_t(\beta)$; * denotes significance of β at the 1 percent level (standard errors have been adjusted for heteroskedasticity and autocorrelation using the Newey and West 1987 estimator); R^2 is the coefficient of determination for the regression. All statistics are estimated for thirty-four observations from 1997:Q1 to 2005:Q2—the longest period for which all data are available.

unobserved target variable (Harding and Pagan 2002). For all of the indicators, concordance is above 50 percent—the proportion of time that a coin toss would accurately predict the correct change in the target variable. Concordance is highest for our core inflation indicator, at 0.79 percent. The exponentially smoothed measure also predicts changes in the target 70 percent of the time.

Following Cogley (2002), the predictive ability of core inflation can be evaluated using the following regression:

$$\pi_{t+4} - \pi_t = \alpha + \beta\left(\pi_t^{core} - \pi_t\right) + \epsilon_{t+4}, \tag{9}$$

where π_t is headline CPI inflation, π_{t+4} is headline CPI inflation one year into the future, π_t^{core} is the core inflation measure, and ϵ_{t+4} is idiosyncratic noise. The regression estimates whether the current gap between core inflation and headline inflation predicts future changes in headline inflation, where predictive power is indicated by $\beta > 0$ (α should also be equal to 0).

We find that our core indicator and the exponentially smoothed indicator have significant predictive power, and they explain nontrivial amounts of the variation of future changes in headline inflation, according to R^2. The other core measures, in contrast, perform very poorly by this metric.

5.1 The Real-Time Properties of the Core Indicator

Because our core indicator is estimated, unlike the other core inflation measures displayed in table 2, it is subject to revision as more data become available. However, this may not impact dramatically on the relative predictive performance of the indicator.

To examine the real-time properties of the indicator, we estimate it recursively for each quarter from 1997:Q1 to 2005:Q2 (the next section describes the real-time estimation procedure in more detail). An indicator of core inflation that is not subject to revision can then be created using the real-time estimates of core inflation in each quarter; i.e., the series $\pi_{t,real}^{core}$ can be compiled using π_t^{core} estimated each quarter from 1997:Q1 to 2005:Q2.

Comparing these real-time estimates of core inflation with core inflation estimated using all of the data (the last row of table 2), we find that the indicator retains many desirable features. It remains

Table 3. Revisions to the Core Indicator in Real Time

	t	$t-1$	$t-2$	$t-3$	$t-4$
Mean Revision	0.12	0.09	0.08	0.07	0.06
Mean Absolute Revision	0.26	0.17	0.10	0.08	0.07

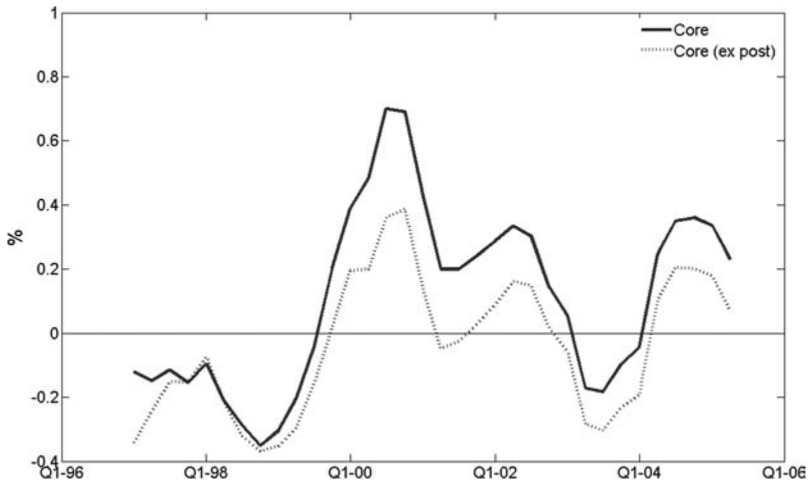
highly correlated with the target, albeit now with a lag of one quarter. Interestingly, the concordance with the target increases with the real-time core indicator.

Revisions to the core inflation indicator over time can be analyzed by comparing the real-time estimates of core inflation with the core inflation indicator estimated using all of the data. To be concrete, each quarter t, the estimate of core inflation can change not only for period t but also for all ω historical periods prior to period t, i.e., periods $t-1,\ldots,t-\omega$, where ω denotes the first observation used to estimate the indicator each quarter. Table 3 displays the mean revisions and the mean absolute revisions to the core indicator at periods $t, t-1,\ldots,t-4$ (revisions are defined to be the core indicator estimated up to the end of the sample minus the real-time estimates).

Not surprisingly, the revisions are larger for the most recent estimates of core inflation. The final five estimates of core inflation have generally been revised up since 1997:Q1, with the revisions to the most recent estimate, period t, being just under twice as large as those from a year prior to the most recent estimate. The magnitude of each revision in period t is 0.26 percent, almost four times more than the estimate from period $t-4$.

Notwithstanding any changes to the dynamic relationships within our panel, with such a small sample to begin with (twenty-one observations), the core inflation will suffer from biases relating to changes to time-series properties of headline CPI inflation over the sample. This can be seen in figure 3, where we see that the mean and variance of the series are higher in the second half of the sample, perhaps as a result of the changes to the inflation target of the Reserve Bank of New Zealand. Recall that, after the dynamic factor model is estimated, the mean and the standard deviation of headline CPI inflation are reattributed to π_t^{core} to make it comparable to

Figure 4. Revisions to Real-Time Estimates of Core Inflation



headline CPI inflation. Mean and variance shifts are thus a source of revision to our core indicator.

To further examine the impact of this type of revision, we compare the real-time estimates of the core inflation, Core, with the real-time estimates of core inflation scaled with the mean and standard deviation of headline inflation estimated over the entire sample, Core (ex post). In this way, we get a sense of how mean and variance shifts over the sample have influenced the size of the revisions. The revisions to the most current estimate of core inflation in real time (period t) from these two core indicators are displayed in figure 4.

Revisions to the standard core indicator were particularly large over 2000, when there was a substantial shift in headline inflation, although the impact of that shift did not influence Core (ex post) by as much. Indeed, generally speaking, Core was not revised as much as Core (ex post), suggesting that mean and variance shifts over our sample period were partly to blame for the revisions to the real-time estimates of core inflation. However, this is only one source of revision to our core indicator. There are many others not considered here, including changes to the dynamic structure of the panel over the sample period. Notwithstanding any changes of this type, it is reasonable to expect better estimates of the mean and standard deviation with which we scale the core inflation indicator as more data come to hand, which suggests that revisions to the indicator will likely be smaller in the future.

We have seen that, regardless of being subject to (sometimes substantial) revision, the core indicator compares favorably with a range of other, more standard, measures of core inflation in predicting an unobserved centered two-year moving average of headline CPI inflation—an approximation of the medium-term inflation target of the RBNZ. In the next section, we further examine the real-time properties of the inflation indicator by way of a real-time prediction experiment for the unobserved target variable.

6. A Real-Time Prediction Experiment

To simulate the predictive performance of our indicator, we estimate the dynamic factor model each quarter from T_0 = 1999:Q4 to T_1 = 2006:Q2, using exactly the information that was available in real time. In New Zealand, the CPI data are not subject to revision and do not require seasonal adjustment since they are expressed in log seasonal differences. The data are outlier-adjusted and standardized prior to estimation in each quarter.

The predictive ability of our indicator is compared with two broad categories of estimators of core inflation. The first category contains methods of estimating core inflation that utilize only CPI data: dynamic factor forecasts, time-series forecasts, conventional core inflation indicators, and static factor model forecasts. The other category contains indicators based on the real-time forecasts from a suite of models used in the policy process at the RBNZ, many of which use a much broader data set than is used to compute our core inflation indicator.

By definition, our core inflation indicator produces an estimate of the unobservable centered moving average of annual inflation. Some of the other methods of estimating core inflation we consider, however, are not tailored to this representation of inflation; forecast-based methods, e.g., do not yield estimates of inflation that are compatible with our target measure. To address this problem, we adopt a forecast-based measure of the target, which averages historical inflation and forecasts of inflation. Specifically, each quarter, inflation is forecast four periods ahead, and an estimate of the unobserved inflation target, π_t^{target}, is computed as a centered moving average of the historical and forecast data. Specifically, the forecast-based estimate of the inflation target based on model i is

$$\hat{\pi}_t^{target}(i) = \frac{1}{9}(\pi_{t-4} + \pi_{t-3} + \ldots + \pi_t + \ldots + \hat{\pi}_{t+3}(i) + \hat{\pi}_{t+4}(i)), \quad (10)$$

where all inflation observations after period t are forecasts from model i.

The estimate of the target is then compared to the ex post target, π_t^{target}; the number of prediction errors is $T_1 - T_0 - 4$. We compute each indicator's root mean-squared error (RMSE) and compare these statistics with the RMSEs from a range of other real-time estimates of the target variable. The RMSE of model i is defined as

$$RMSE(i) = \sqrt{\frac{1}{T_1 - T_0 - 4} \sum_{t=T_0}^{T_1-4} \left(\pi_t^{target} - \hat{\pi}_t^{target}(i)\right)^2}. \quad (11)$$

We use the Diebold and Mariano (1995) statistic to assess the quality of our results, testing whether the MSEs (the square of the RMSE) of the competing models are statically different from the dynamic factor model estimate of the target. The test statistic for model i is

$$STAT(i) = \frac{\bar{d}(i)}{\sqrt{\hat{V}(\bar{d}(i))}}, \quad (12)$$

where $\bar{d}(i)$ is the mean difference in squared errors $(e_t^2(i) - e_t^2)$ between model i and the dynamic factor model estimate of the target, $e_t(i)$ is the forecast error from model i, and e_t is the forecast error from the dynamic factor model.[14]

To test whether the forecasts make a useful contribution to an estimate of the inflation target estimated using the published forecasts of the RBNZ (discussed below), we also compute a variant of

[14]In practice, $STAT(i)$ is tested using a t-test with $T_1 - T_0 - 4$ degrees of freedom. The variance of the error differentials $V(\bar{d}(i))$ is adjusted for heteroskedasticity and autocorrelation using the Newey and West (1987) estimator, with a truncation lag of 3.

the Chong and Hendry (1986) encompassing test. This is based on the following forecast combination regression:

$$\pi_t^{target} = \lambda \hat{\pi}_t^{target}(i) + (1-\lambda)\hat{\pi}_t^{target}(RBNZ) + \epsilon_t, \tag{13}$$

where $\hat{\pi}_t^{target}(i)$ is the prediction from model i, $\hat{\pi}_t^{target}(RBNZ)$ is the prediction from RBNZ's published forecasts, and ϵ_t is an idiosyncratic error term. If $\lambda = 0$, model i adds nothing to the RBNZ predictions, and if $\lambda = 1$, the RBNZ predictions add nothing to the predictions from model i.[15]

We define the following forecasting model for annual headline CPI inflation:

$$\pi_{t+h} = \beta_0 + \sum_{j=0}^{p} \beta_{1j}\pi_{t-j} + \sum_{j=0}^{k}\sum_{m=1}^{r} \beta_{2jm} x_{m,t-j} + \epsilon_{t+h}, \tag{14}$$

where π_{t+h} is inflation h periods into the future ($h = 1, \ldots, 4$), π_t is inflation in period t, and $x_{m,t}$ is a variable used to predict inflation.

6.1 The Models Based on Disaggregate CPI Series

6.1.1 Core Inflation Indicator

The dynamic factor model is estimated with $q = 4$ and $s = 2$, intertemporally smoothing the common component by summing waves of frequency between $[-\pi/9, \pi/9]$. The estimate of the target is simply $\hat{\chi}_t^L$.[16]

6.1.2 Dynamic Factor Model Forecasts

As with the core indicator, the dynamic factor model is estimated with $q = 4$ and $s = 2$. Here, however, we forecast the common component $\hat{\chi}_t$ itself (Forni et al. 2005 show how this done in practice—a brief summary of the procedure can be found in appendix 1).

[15] As with the Diebold and Mariano (1995) test, the standard errors of the regression are adjusted using the Newey and West (1987) estimator.

[16] Appendix 4 displays the RMSEs for different configurations of q and s.

These forecasts are concatenated with historical inflation data and averaged using equation (10) to produce $\hat{\pi}_t^{target}$.[17]

6.1.3 Standard Core Inflation Indicators

We compute two estimates of the inflation target using a range of standard core inflation indicators: trimmed-mean inflation; weighted-median inflation; the CPI excluding food, administration charges, and gasoline; median inflation; double-weighted inflation (Wynne 1997) ; and exponentially smoothed inflation (Cogley 2002). Appendix 2 describes these indicators in greater detail. The first estimate of the target is the "naive" prediction based on each indicator's raw estimate of core inflation, $\hat{\pi}_t^{target} = \pi_t^{core}$. The second estimate uses equation (14) to forecast inflation four quarters into the future. Specifically, in (14), $m = 1$ and $x_{m,t} = \pi_t^{core}$. In each quarter, and at each forecasting horizon, the lag orders p and k are selected using the Schwartz-Bayesian information criteria (BIC), where $p = 0, \ldots, 3$ and $k = 0, \ldots, 3$ (so that the maximum number of lags of each variable is four).[18] These forecasts are then used to compute $\hat{\pi}_t^{target}$, based on equation (10). This is called the "scaled" standard core inflation indicator.

6.1.4 Time-Series Forecasts

Three time-series models are used to forecast headline inflation one year ahead: autoregressive, random walk, and random walk in mean. In each quarter, and at each forecasting horizon, the autoregressive forecast is made using equation (14), where $\beta_{2jm} = 0$, and the autoregressive order p is chosen using the BIC, with $p = 0, \ldots, 3$.

[17]A core inflation indicator similar to this produced the best forecasts of annual inflation in Cristadoro et al. (2005). Cristadoro et al. (2005), however, concatenate their dynamic factor model forecasts with the in-sample predicted values from the dynamic factor model. We chose to concatenate with historical data instead, because the resulting indicator produced slightly better predictive performance.

[18]The BIC for model i and horizon h is defined as $BIC^h(i) = T\ln(S/T) + 2k\ln(T)$, where T is the number of usable time-series observations, k is the number of estimated parameters, and S is the sum of squared errors of the regression, $S = \sum_{t=1}^{T}(\pi_{t+h} - \hat{\pi}_{t+h})^2$.

The random-walk forecast is made using equation (14) with $\beta_0 = 0$, $p = 0$, $\beta_{10} = 1$, and $\beta_{2jm} = 0$, and the random-walk-in-mean forecast takes the mean of the series as the forecast at all horizons (all terms in equation (14) are set to 0 except for β_0). Once the forecasts are made, they are concatenated with historical data to yield a prediction for the inflation target, based on equation (10).

6.1.5 Pooling Regressions

We make inflation forecasts from one to four quarters ahead using equation (14) and each of the i components of the CPI, where $m = 1$, $x_{m,t} = \pi_{i,t}$, and $i = 1, \ldots, n$. The lag orders p and k are selected using the BIC at each horizon h, where $p = 0, \ldots, 3$ and $k = 0, \ldots, 3$. The resulting 227 forecasts for $h = 1, \ldots, 4$ are then weighted using three methods: a simple average (equal weights), a BIC-weighted average, and an expenditure-weighted average. To be explicit, at each forecasting horizon, let the weighted-average forecast for headline CPI inflation be

$$\hat{\pi}_{t+h} = \sum_{i=1}^{n} \Omega^h(i)\hat{\pi}_{t+h}(i), \tag{15}$$

where $\Omega^h(i)$ is the weight attached to model i's forecast at horizon h and $\hat{\pi}_{t+h}(i)$ is the forecast from model i at horizon h. The average forecast sets the weights equal across the n forecasts so that $\Omega^h(i) = 1/n$ for all h. The BIC-weighted average weights each forecast according to the fit that the underlying estimated model had to the data, weighting models with better fit more highly. Here, the weight attached to each forecast i at each horizon h is calculated as

$$\Omega^h(i) = \frac{\exp(-0.5BIC^h(i))}{\sum_{j=1}^{n} \exp(-0.5BIC^h(j))}, \tag{16}$$

where $BIC^h(i)$ is the minimum BIC for model i at horizon h over models estimated with $p = 0, \ldots, 3$ and $k = 0, \ldots, 3$.[19]

[19]The BIC-weighted-average forecast is approximately the Bayesian-model average forecast arising from equal model priors and diffuse coefficient priors.

The expenditure-weighted average weights the forecast associated with CPI component i with that component's expenditure weight, w_i, across all horizons h, $\Omega^h(i) = w_i$.

The three sets of pooled forecasts are concatenated with historical data and averaged using equation (10) to yield predictions for the target measure of inflation.

6.1.6 Static Factor Model Forecasts

The static factor model forecasts for headline inflation are made using the first r static principal components estimated from the panel of CPI data, $f_{1,t}, \ldots, f_{r,t}$. For each quarter, and each forecasting horizon h, equation (14) is estimated with $x_{m,t} = f_{m,t}$, where the orders of p, k, and r are selected using the BIC, with $r = 1, \ldots, 4$; $p = 0, \ldots, 3$; and $k = 0, \ldots, 3$. These forecasts are then concatenated with historical data and averaged using equation (10).

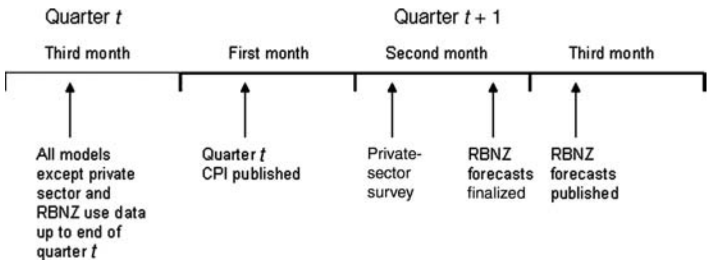
6.2 Core Indicators Based on Some of the Forecasting Models Used at the RBNZ

A prediction of the inflation target at time t is computed using these forecasts in the same way as for the time-series forecasts from above: the real-time estimate of core inflation at time t is $\hat{\pi}_t^{target}$, constructed using equation (10), where all observations up to and including period t are actual data and all observations beyond period t are forecasts. Some of these models use a much broader information set than the estimates of the target rate of inflation described above (see appendix 3 for a detailed description of these forecasts). The forecasting models are as follows:

- the real-time published forecasts of the RBNZ
- an average of real-time private-sector forecasts
- a Bayesian VAR forecast (BVAR)
- a BIC-weighted VAR forecast
- a factor-model forecast
- an indicator forecast

The published forecasts and the average of private-sector forecasts require further discussion, as they can be characterized as

Figure 5. Calender of CPI Publication and Forecasts



being judgment-based forecasts—unlike the other forecasts discussed so far.

With the exception of the core indicators based on these two forecasts, which can incorporate information dated in period $t+1$, all other core indicators discussed so far use information dated up to the end of period t—the same period for which the latest CPI data are available. This can be better illustrated in figure 5. The majority of the estimators of core inflation are made with information up to the end of quarter t; estimates of $\hat{\pi}_t^{target}$ can be constructed using most models with the arrival of period t's CPI data, published in the middle of the first month of period $t+1$. However, the RBNZ and private-sector forecasts incorporate information up to the second month of period $t+1$, using more up-to-date information than the other models.

7. Empirical Results

The results from the real-time prediction experiment are displayed in table 4.

Looking first at the indicators based on the disaggregate CPI data, the core indicator and the scaled exponentially smoothed indicator compare favorably with the other estimates of core inflation, producing the lowest RMSEs (around 0.26 percent) of the price-based indicators examined. The other indicators perform much worse than these two indicators, with several indicators having predictive performance significantly worse than the core indicator (at the 10 percent level). Aside from the core indicator and the exponentially smoothed indicator, the standard core inflation indicators perform

Table 4. Prediction Statistics

Indicators Based on Disaggregate Prices		RMSE		Encompassing (λ)	
Dynamic Factor Models	Core Indicator		0.257		0.270*
	Forecast		0.310		0.114*
Time-Series Models	AR		0.323*		0.087*
	Random Walk		0.342		0.051
	Random Walk (Mean)		0.342		0.108*
		Naive	*Scaled*	*Naive*	*Scaled*
Core Inflation Indicators	Weighted Median	0.630*	0.390	−0.052	−0.058
	Trimmed Mean	0.500*	0.791	−0.049	−0.033
	CPI exl. Food, Admin., and Gasoline	1.011*	0.358	−0.090	−0.004
	Median	0.640*	0.408*	−0.015	0.049
	Double Weighted	0.399	0.328	−0.033	0.060
	Exponentially Smoothed	0.310	0.258	0.146*	0.269*
Pooling Regressions	Average		0.325*		0.085*
	BIC Weighted		0.325*		0.085*
	Expenditure Weighted		0.332*		0.067
Static Factor Model	Static Factor Model		0.337		0.034
Indicators Based on RBNZ Forecasts					
RBNZ Models	Published		0.186		—
	External Average		0.123*		1.002*
	BVAR		0.291		0.065
	VAR		0.290		0.165*
	Factor Model		0.264		0.123
	Indicator Median		0.259		0.165*

Note: * denotes statistical significance at the 10 percent level. RMSE: A Diebold and Mariano (1995) test is used, testing whether the MSE of model i is statistically different from the core indicator. Encompassing (λ): A Chong and Hendry (1986) test is used, testing whether there is predictive content in model i over and above the predictive content of the indicator based on the published forecasts of the RBNZ, $\lambda \neq 0$.

particularly poorly, though (generally speaking) the performance of these indicators improves when scaled (used in a forecasting regression). The remainder of the models based on the disaggregate CPI data have comparable predictive ability, each with RMSEs of between 0.30 and 0.35 percent.

Overall, the indicators based on the published forecasts of the RBNZ and the private-sector forecasts are the best predictors of the target variable, followed by the core inflation indicator and the scaled exponentially smoothed indicator: the indicator based on the private-sector forecasts is particularly good and statistically better than the core inflation indicator. The relatively good performance of the RBNZ and private-sector indicators comes as no surprise, given that they are based on an information set that is both broader and more up-to-date than is used by the other indicators. Indeed, aside from the core inflation indicator and the scaled exponentially smoothed indicator, the indicators based on disaggregate CPI data are worse than the indicators based on RBNZ forecasts, perhaps reflecting the broader range of information that is incorporated in these indicators (e.g., real variables are incorporated into all RBNZ models).

Despite being altogether worse than the published forecasts of the RBNZ according to RMSE, some of the core indicators are able to improve the predictive performance of the RBNZ indicator. Looking at weights attached to each of the indicators relative to the RBNZ indicator from the encompassing regression λ, we find that the largest weights are attached to the external average, the core indicator, and the scaled exponentially smoothed indicator. Moreover, because the core inflation indicator and the scaled exponentially smoothed indicator can be computed as soon as the CPI data are available, these indicators are more timely than the indicator based on published RBNZ forecasts, which are finalized more than one month later.

8. Summary and Conclusion

This paper introduced a new indicator of core inflation for New Zealand, estimated using a dynamic factor model. Defining core inflation to be consistent with the medium-term inflation target of the RBNZ, we found that our indicator produced relatively accurate estimates of core inflation, compared with a range of other indicators of core inflation. Estimates of core inflation derived from the RBNZ's published forecasts and an average of private-sector forecasts produce more accurate measures of core inflation than our indicator. However, our core indicator is more timely and can be computed as soon as the CPI data are published—around a month before the RBNZ forecasts are finalized.

Appendix 1. Estimation of Static and Dynamic Factor Models

Static Factor Model

Assume there are T time-series observations for N cross-section units denoted x_{it}, where $i = 1, \ldots, N$ and $t = 1, \ldots, T$. We let X be the $T \times N$ matrix of observations, x_t is a row denoting all N observations at time t, and x_j is a column vector denoting all T observations for cross-section unit j.

The static factor model is estimated using an eigenvector-eigenvalue decomposition of the sample covariance matrix—principal components. Specifically, let V be the eigenvectors corresponding to the r largest eigenvalues of the $N \times N$ matrix $\hat{\Gamma} = \frac{1}{T}\sum_{t=1}^{T} x_t x_t'$. The static principal components estimator yields

$$\hat{F} = XV, \quad \hat{\Lambda} = V, \quad \text{and} \quad \hat{\chi} = \hat{F}\hat{\Lambda}', \tag{17}$$

where $\hat{F}$ is a $T \times r$ matrix of common factors, $\hat{\Lambda}$ is an $r \times N$ matrix of factor loadings, and $\hat{\chi}$ is a $T \times N$ matrix of common components.

Dynamic Factor Model

The dynamic factor model is estimated using an eigenvalue decomposition of the spectrum smoothed over a range of frequencies—dynamic principal components. Estimation proceeds as follows:

(i) Estimate the spectral-density matrix of X, using a Bartlett lag window of size M; i.e., compute the autocovariance matrices $\hat{\Gamma}_X(k) = \frac{1}{T}\sum_{t=k+1}^{T} x' x_{t-k}$, where $k = 1, \ldots, M$, multiply them by the weights $\omega_k = 1 - \frac{|k|}{M+1}$, and apply the discrete Fourier transform:

$$\hat{\Sigma}_X(\omega_m) = \frac{1}{2\pi} \sum_{k=-M}^{M} \omega_k \Gamma(k) e^{-i\omega_m k}. \tag{18}$$

(ii) For each frequency, $\omega_m = \frac{2\pi m}{2M+1}$, $m = -M, \ldots, M$, let $D_q(\omega_m)$ be the diagonal matrix with the q largest eigenvalues of $\hat{\Sigma}_X(\omega_m)$ on the diagonal, and let $U_q(\omega_m)$ be the associated matrix of eigenvectors. Use the discrete inverse Fourier transform on $\hat{\Sigma}_\chi(\omega_m) = U_q(\omega_m) D_q(\omega_m) U_q(\omega_m)'$ to obtain

$$\hat{\Gamma}_\chi(k) = \frac{2\pi}{2M+1} \sum_{m=-M}^{M} \hat{\Sigma}_\chi(\omega_m) e^{i\omega_m k}. \tag{19}$$

(iii) The covariance matrix of the idiosyncratic part is then estimated as a residual:

$$\hat{\Gamma}_\epsilon(k) = \hat{\Gamma}_X(k) - \hat{\Gamma}_\chi(k). \tag{20}$$

(iv) Let Z be the r generalized eigenvectors (with eigenvalues in descending order) of $\hat{\Gamma}_\chi(0)$ with respect to $\hat{\Gamma}_\epsilon(0)$ with the normalization that $Z_j \hat{\Gamma}_\epsilon(0) Z_i' = 1$ if $i = j$ and 0 otherwise. This is generalized principal components, which is essentially weighted principal components, where each principal component is weighted by the inverse of the size of its idiosyncratic noise.

(v) The estimated dynamic factors and common components are, respectively,

$$\hat{F} = XZ \quad \text{and} \quad \hat{\chi}_{t+h} = \hat{\Gamma}_\chi(h) Z (Z' \hat{\Gamma}_X(0) Z)^{-1} Z' x_t \tag{21}$$

for $h = 0, \ldots, M$.

The covariance matrix of the long-run common components, $\hat{\Gamma}_\chi^L(\omega_m)$, can be estimated by applying the inverse Fourier transform (step ii above) to the spectral-density matrix (18) over the frequency band of interest $[-\pi/\tau, \pi/\tau]$ (where τ denotes the periodicity of the shortest cycle allowed). The estimated long-run common components are then

$$\tilde{\chi}^L_{t+h} = \hat{\Gamma}^L_{\chi}(h)Z(Z'\hat{\Gamma}_X(0)Z)^{-1}Z'x_t. \tag{22}$$

Appendix 2. The Standard Core Inflation Measures

Let headline CPI inflation be the weighted average of the inflation rates of n disaggregate subgroups of the CPI:

$$\pi_t = \sum_{i=1}^{n} w_{it}\pi_{it}, \tag{23}$$

where π_t is headline CPI inflation, π_{it} is inflation in the i-th component of the CPI, and w_{it} is the expenditure weight of the i-th component.

Trimmed Mean

Each quarter, the trimmed mean is calculated as follows:

(i) Compute the annual percentage change in each of the i components of the CPI.

(ii) Sort the resulting series from smallest to largest, along with their associated weights w_i.

(iii) Compute the cumulative sum of the weights of the ordered prices.

(iv) Exclude the series with cumulative weights either less than 5 percent or with cumulative weights greater than 95 percent.

(v) Compute the trimmed-mean inflation rate as

$$[1/\sum_{i=first}^{last} w_i] \sum_{i=first}^{last} w_i\pi_{it}, \tag{24}$$

where *first* and *last* denote the first and last CPI components in the truncated list of ordered price changes.

Weighted Median

The weighted-median inflation rate is calculated using steps (i)–(iii) above. The weighted median is then the first percentage change in price where the cumulative weight is greater than or equal to 50 percent.

Median

Median inflation is simply the median rate of inflation from the n components of the CPI.

CPI Excluding Food, Administration Charges, and Gasoline

This measure is computed using

$$\sum_{i=\Omega}^{n} w_i \pi_{it}, \tag{25}$$

where Ω is the number of CPI components categorized as food, administration charges, and gasoline, and the weights w_i have been adjusted to sum to 1.

Double-Weighted Inflation

This measure of core inflation is computed by each of the components of the CPI—a weight inversely proportional to its variability. Double-weighted inflation is calculated as

$$\frac{\sum_{i=1}^{n} w_i v_i \pi_{it}}{\sum_{i=1}^{n} w_i v_i}, \quad \text{where} \quad v_i = \frac{1/\sigma_{it}}{\sum_{i=1}^{n} 1/\sigma_{it}}, \tag{26}$$

and where σ_{it} is the standard deviation of inflation in component i relative to headline CPI inflation, $(\pi_{it} - \pi_t)$.

Exponentially Smoothed Inflation

Exponentially smoothed inflation is

$$\phi \sum_{i=1}^{j} (1-\phi)^j \pi_{t-i}, \tag{27}$$

where ϕ, the expectations adjustment parameter, is set to 0.125, as in Cogley (2002).

Appendix 3. Some Forecasts Used at the RBNZ

The Published Forecasts

These are the real-time forecasts published in the Reserve Bank's quarterly *Monetary Policy Statement* (*MPS*). The forecasts are a combination of model-based forecasts and judgment. Broadly speaking, the near-term forecasts can be characterized as being judgment based and indicator based. The longer-term forecasts, on the other hand, are made with the help of a large-scale macroeconomic model, the Forecasting and Policy System (FPS).

External Average

The external-average forecast is the average forecast from a group of private- and public-sector institutions. These forecasts are attained by an informal survey conducted by the RBNZ in the middle of each quarter.

BVAR

The BVAR has a Minnesota prior, shrinking the VAR coefficients toward univariate unit roots (Doan, Litterman, and Sims 1984), and contains five endogenous variables n (GDP, CPI, ninety-day rates, the trade-weighted index [TWI], and the terms of trade [TOT]). The VAR can be written as

$$y_t = \Phi_0 + \Phi_1 y_{t-1} + \ldots + \Phi_p y_{p-t} + u_t, \tag{28}$$

where u_t is a vector of one-step-ahead forecast errors and y_t is $n \times 1$. We assume that the innovations in (28) have a multivariate normal distribution $N(0, \Sigma_u)$. The VAR can also be expressed in matrix form as

$$Y = X\Phi + U, \tag{29}$$

where Y is a $T \times n$ matrix with rows y_t'; X is a $T \times k$ matrix with rows $x_t' = [1, y_{t-1}', \ldots, y_{t-p}']$, where $k = 1 + np$; U is a $T \times n$ matrix with rows u_t'; and $\Phi = [\Phi_0, \ldots, \Phi_p]'$. The Minnesota prior is implemented as in Del Negro and Schorfheide (2004):

$$\bar{\Phi} = (I \otimes (X'X) + \iota H^{-1})^{-1}(\text{vec}(X'Y) + \iota H^{-1}\Phi), \tag{30}$$

where the ι denotes the weight of the Minnesota prior, Φ is the prior mean, and H is the prior tightness. The values for Φ and H are the same as in Doan, Litterman, and Sims (1984). All variables, except ninety-day interest rates, enter the VAR in log-differences. Thus, to be consistent with the Minnesota prior, the prior for the mean of the first lag of all growth variables is 0; the prior for the mean of the first lag of the ninety-day interest rate is 1. The prior is augmented with a proper inverse Wishart prior for Σ_u. The overall tightness of the prior is determined by the hyperparameter ι. Following Del Negro and Schorfheide (2004), this parameter is chosen ex ante to maximize the marginal data density; i.e., each quarter,

$$\hat{\iota} = \arg\max_{\iota} p_{\iota}(Y). \tag{31}$$

VAR

The VAR forecast is a BIC-weighted average of VAR forecasts from over 500 different VAR specifications. The VAR specifications differ in three respects: (i) they have different variables in them, (ii) they embody a different number of lags (between one and three lags for each variable), and (iii) some of the VARs are in difference form while others use levels data. The base VAR uses GDP, CPI, the ninety-day bank-bill rate and the (real/nominal) TWI. This base model is then augmented with a world sector, commodity prices, migration and housing, and hours worked. The world sector is sometimes represented using U.S. GDP, U.S. CPI, and U.S. interest rates, and sometimes using the inputs that feed into FPS—world GDP (defined to be a twelve-country weighted average), short world interest rates, and world CPI. The commodity prices also take various forms: the ANZ SDR commodity price index, (import and) export prices denominated in New Zealand dollars, and a U.S.-dollar oil price. Because of limitations in the data, not all of these series can be incorporated in the model simultaneously, which is why the VAR forecast is obtained over an average of models.

Factor Model

The factor-model forecast is derived from a data set comprising almost 400 macroeconomic series. All series in the data set are seasonally adjusted using X12 (additive). The series are then transformed to account for stochastic and deterministic trends; the I(1) series are logged and then differenced, and the I(0) series are left as levels. See Matheson (2006) for a more detailed description of these data.

We estimate static factors from this data set using the same method as described above for the static factor models and use them in the following h-step-ahead regression:

$$\pi_{t+h} = \phi + \beta(L)f_t + \gamma(L)\pi_t + e_{t+h}, \tag{32}$$

where $\pi_{t+h} = \ln(p_{t+h}/p_t)$ is h-period inflation in CPI P_t and $\pi_t = \ln(P_t/P_{t-1})$, ϕ is a constant, $\beta(L)$ and $\gamma(L)$ are lag polynomials, f_t is a vector of factors (estimated using static principal components), and e_{t+h} is an error term.

The algorithm that is used to produce factor-model forecasts each quarter tailors the raw data set X to the task of forecasting inflation at different horizons: note that the factors from X are the same regardless of the horizon being forecast. Following the method described in Matheson (2006), the data set X is reduced by removing those series that do not have a high correlation with inflation at the horizon being forecast. Essentially, we regress the inflation rate that we are trying to forecast on each series in the data set. We then rank the resulting R-squareds (coefficients of determination) and remove those series that are least informative—keeping a proportion θ of the series at each horizon h. The factors are then extracted from this reduced data set X^*.

Due to uncertainty about the particular cut-off criterion to choose, we average the factor-model forecasts over a variety of different criteria: $\theta = (5, 10, 20, 50, 100)$.[20] Aside from the size of the data set, the five factor-model forecasts are constructed in the same way using (32). We estimate (32) using the factor with the largest

[20]Note that when $\theta = 100$ all of the data are used to extract factors, as in Stock and Watson (2002).

eigenvalue (the principal component) and do not allow lags of the factor ($\beta(L) = \beta$). The number of lags of inflation that are included at each horizon is chosen using the BIC, with lags varying from zero to four.

Indicator Forecast

The indicator median forecast uses the same data set and forecasting methodology as the factor-model forecast. Bivariate regressions are run for each series in the data set, where series x_i replaces f_t in (32). The number of lags of each indicator is allowed to vary from one to four and the number of lags of inflation is allowed to vary from zero to four, with all lags selected with the BIC. The BICs from these regressions are then ranked. The indicator median forecast is the median forecast from the top 10 percent of the ranked bivariate regressions.

Appendix 4. Core Indicator: RMSE for Different Configurations of *q* and *s*

	s	**0**	**1**	**2**	**3**	**4**
q						
1		0.465	0.493	0.461	0.471	0.478
2		0.432	0.370	0.364	0.362	0.368
3		0.340	0.320	0.327	0.319	0.320
4		0.241	0.257	***0.257***	0.274	0.290
5		0.235	0.265	0.279	0.293	0.294
6		0.254	0.264	0.290	0.300	0.302

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Firm-Specific or Household-Specific Sticky Wages in the New Keynesian Model?*

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This paper shows that switching the dominant use of household-specific sticky wages in the New Keynesian model (Erceg, Henderson, and Levin 2000) for firm-specific sticky wages has qualitative and quantitative consequences. First, the model with firm-specific sticky wages incorporates endogenous changes in the rate of unemployment, whereas there is no unemployment with household-specific sticky wages. Secondly, business-cycle fluctuations of wage inflation and the real wage are clearly distinguishable. In particular, the real wage is countercyclical after a demand shock under any sensible calibration with firm-specific sticky wages, whereas the model with household-specific sticky wages requires larger wage stickiness than price stickiness. Finally, optimal monetary policy is more oriented to stabilizing price inflation with firm-specific sticky wages, and is more oriented to stabilizing the output gap and wage inflation with household-specific sticky wages.

JEL Codes: E12, E24, E32, J30.

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1. Introduction

The New Keynesian framework was initially introduced as a general-equilibrium (microfounded) structure with nominal rigidities on price setting (King and Wolman 1996; Yun 1996). There was a competitive labor market with homogeneous labor and market-clearing wages. Soon it was noticed that a flexible-wage labor market, even combined with sticky prices, brings over some business-cycle patterns that are difficult to find in actual data—mainly, high volatility on nominal and real wages, strongly procyclical real wages, and little persistence on price inflation and output.[1] To dampen wage volatility, most early New Keynesian models were specified with a high labor-supply elasticity to the real wage (as usually assumed in the real-business-cycle [RBC] literature with logarithmic specifications for the leisure component in the utility function). However, the bulk of empirical microevidence suggests that the labor-supply elasticity should be a low positive number (Altonji 1986; Pencavel 1986; and Card 1994), disputing the macrolevel calibration commonly used in the New Keynesian literature.

Staggered-wage contracts may also help to reduce wage volatility.[2] In a well-known paper, Erceg, Henderson, and Levin (2000)—EHL henceforth—describe a New Keynesian model with both staggered prices and staggered-wage contracts that can be optimally adjusted subject to some constant probability à la Calvo (1983).[3,4]

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[1]See Jeanne (1998), Taylor (1999), Chari, Kehoe, and McGrattan (2000), Casares (2002), and Krause and Lubik (2007).

[2]Other sources of wage rigidities, such as efficiency wages (Danthine and Kurmann 2004) or search and matching frictions (Krause and Lubik 2007), also have been recently incorporated into the labor market of the New Keynesian model.

[3]The Calvo (1983) constant-probability assumption is frequently taken in New Keynesian models because it results in a simple and comprehensive inflation equation: the so-called New Keynesian Phillips curve (Walsh 2003, chap. 5). In turn, inflation fluctuates in the New Keynesian model, responding to changes in current and expected future real marginal costs.

[4]Other authors use Taylor (1980) staggered-price contracts with a predetermined length (Chari, Kehoe, and McGrattan 2000; Huang, Liu, and Phaneuf 2004), and some others follow Rotemberg (1982) to assume a quadratic price-adjustment cost function to slow down price adjustments (Ireland 2003). In all cases, the resulting price-inflation dynamics are similar to those obtained with the dominant Calvo-style pricing scheme.

Households act as wage setters and firms as price setters.[5] Thus, each household owns one differentiated type of labor service and can decide on its nominal-wage rate, provided the arrival of the adequate Calvo-type market signal. As a result, the dynamics of wage inflation can be formulated in a single forward-looking equation governed by the gap between the aggregate marginal rate of substitution of households and the real wage. The EHL model with household-specific sticky wages has received empirical support (Galí, Gertler, and López-Salido 2001; Rabanal and Rubio-Ramírez 2005) and is becoming a preeminent model for monetary policy analysis (Amato and Laubach 2003; Smets and Wouters 2003; Woodford 2003; Giannoni and Woodford 2004; Christiano, Eichenbaum, and Evans 2005; Levin et al. 2006; Casares 2007a).

This paper describes one variant for a sticky-wage New Keynesian model in which firms are the wage-setting actors instead of households. Wage contracts will be reset only in cases when the firm is able to post the optimal price, attaching the Calvo-style staggered-prices scheme also to staggered wages. As in Bénassy (1995), firms will offer households the nominal wage that matches their labor demand with the households' labor supply. These firm-specific sticky wages result in a wage-inflation equation different from that obtained in the EHL model with household-specific sticky wages. In particular, wage-inflation fluctuations are influenced by two real-wage gaps: the household-related gap between the marginal rate of substitution and the real wage (also present in the EHL model) and the firm-related gap between labor productivity and the real wage (absent in the EHL model).

Furthermore, firm-specific sticky wages bring to the New Keynesian model an endogenous measure of unemployment due to the separation between demand and supply of labor in the fraction of wage contracts that cannot be renegotiated over the current period. The unemployment rate is then obtained as the percent difference between economy-wide labor supply and labor demand. In recent years, several papers have already shown how to incorporate unemployment into a New Keynesian structure by attaching a labor

[5] The assumption of providing households with market power to set wages had already been taken in Blanchard (1986), Rankin (1998), and Ascari (2000).

market with matching frictions à la Mortensen-Pissarides (1994).[6] In those papers, unemployment arises as a result of having search costs and matching frictions in the labor market. This paper shows the alternative of nominal rigidities on the firm-specific wage-setting procedure to explain the presence of unemployment in the labor market.[7]

The quantitative implications of having either firms or households as wage-setting actors are examined within a complete New Keynesian model with sticky wages. Thus, the business-cycle properties of the EHL model (households set wages) are compared to those of the model with wage-setting firms. Apart from the key difference on the absence or presence of unemployment, we will show how wage inflation and the real wage respond in a substantially different way to both supply and demand shocks. Special attention will be focused on the real-wage business cycle. Sumner and Silver (1989) provide empirical arguments that explain the slight procyclicality of the real wage observed in the U.S. economy as a combined reaction to supply and demand shocks. They show that the real wage is procyclical in periods dominated by supply shocks, whereas it behaves countercyclically in periods when output fluctuations are driven by demand shocks.[8] The impulse-response functions obtained in the New Keynesian model with firm-specific sticky wages provide the kind of real-wage reactions consistent with the Sumner-Silver hypothesis. This result is not found in the baseline calibration of the EHL model with the same level of price and wage stickiness because the real wage is procyclical after a demand shock. The latter is reversed and the EHL model also replicates the Sumner-Silver hypothesis when wage stickiness is higher than price stickiness to let prices react more strongly than wages and obtain a countercyclical real wage.

[6]A list of those papers should include Christoffel and Linzert (2005), Gertler and Trigari (2006), Krause and Lubik (2007), and Kuester (2007).

[7]Blanchard and Galí (2007) also introduce unemployment in a New Keynesian model by assuming rigidities on the aggregate real wage based on one ad hoc formulation.

[8]Fleischman (1999) corroborates these results in his empirical analysis. Bordo, Erceg, and Evans (2000) use a sticky-wage model to argue that during the Great Depression the real wage moved anticyclically in response to a monetary contraction, which is one example of a demand-side shock.

The consequences of firm-specific or household-specific sticky wages for the optimal design of monetary policy in the New Keynesian model are also discussed in this paper. Following Woodford (2003, chap. 8) and Giannoni and Woodford (2004), optimal monetary policy can be obtained by minimizing a welfare-theoretic intertemporal loss function subject to a set of model equations. We compute the optimal monetary policy in the two sticky-wage variants and compare their stabilizing results. In addition, the performance of an instrument Taylor (1993)-type rule is examined both for a (baseline) standard representation and also for one specification that uses optimized coefficients in accordance with optimal monetary policy. The EHL model implies an optimal policy that stabilizes almost completely the output gap with a very high reaction coefficient in the Taylor-type rule. By contrast, the model with firm-specific sticky wages has an optimal monetary policy with a stronger concern on stabilizing price inflation and more output-gap variability as a result. In both sticky-wage cases, the optimized Taylor-type rule provides a stabilizing performance only slightly inferior to that under the optimal monetary policy.

The remaining sections of the paper are organized as follows. Section 2 describes how to introduce sticky wages that are linked to the staggered-pricing behavior of monopolistically competitive firms and how these firm-specific sticky wages explain the presence of unemployment in the labor market. In section 3, we derive the price-inflation equation with firm-specific sticky wages, which involves terms on expected next period's inflation, the real marginal costs, and the rate of unemployment. The complete New Keynesian models with either firm-specific or household-specific sticky wages are outlined in section 4 with numerical values assigned to the structural parameters mostly borrowed from EHL (2000). Sections 5 and 6 are devoted to carrying out the business-cycle and monetary policy comparisons of the two sticky-wage variants. Section 7 concludes with a review of the major findings and contributions of the paper.

2. Firm-Specific Sticky Wages and Unemployment

Let us begin the analysis by describing the wage-setting process of monopolistically competitive firms á la Dixit and Stiglitz (1977), which may jointly decide the price and the nominal wage. Price stickiness causes output, labor demand, and prices to be firm specific because these variables would depend on when was the last time a firm was able to set the optimal price. In principle, if firms can also decide on the nominal wage, they will post values subordinated to the pricing decision and the upcoming labor-demand constraint.

For explanatory purposes, we can split this connection between prices and wages set at the firm into three separate stages. First, the firm-specific price of some i-th firm, $P_t(i)$, determines the amount of output produced by the firm, $y_t(i)$, at the Dixit-Stiglitz demand curve. The higher the price, the lower the output demand with a constant elasticity. Given a production technology, labor demand, $n_t^d(i)$, is then determined as the amount of work hours that must be employed to produce the given level of output. Labor demand, therefore, increases with output. Finally, the labor supply provides the nominal wage, $W_t(i)$, that the firm must set to convince the household to work the number of hours determined by labor demand. Since labor supply responds positively to a wage increase, the firm will offer higher wages when labor demand is rising. In a schematic way, $P_t(i)$ and $W_t(i)$ are connected through this chain:

$$P_t(i) \overset{\textit{Demand (-)}}{\longrightarrow} y_t(i) \overset{\textit{Technology (+)}}{\longrightarrow} n_t^d(i) \overset{\textit{Labor supply (+)}}{\longrightarrow} W_t(i).$$

As a result, the subordinate wage, $W_t(i)$, takes the value obtained when matching the firm-specific labor demand with the household's labor supply. Therefore, a labor-demand equation, a labor-supply curve, and one equilibrium condition are required for the computation of $W_t(i)$.

We start by describing the labor-supply behavior. Unlike the EHL (2000) model that bears household-specific labor, the assumption of firm-specific wages is consistent with an economy where identical households own all the heterogeneous labor services, while firms only demand one differentiated type of labor.[9] Thus, prices

[9]The assumption of firm-specific labor has been recently introduced in the New Keynesian literature. Thus, Woodford (2003, chap. 3) and Matheron (2006) take this assumption with Calvo staggered prices and fully flexible nominal wages.

and nominal wages are set at the firm, whereas households make optimal substitutions across quantities of differentiated consumption goods and labor services.[10] The following separable utility function ranks preferences between bundles of consumption, c_t, and bundles of labor services supplied, n_t^s, for the representative household:

$$U\left(\chi_t, c_t, n_t^s\right) = \exp(\chi_t)\frac{c_t^{1-\sigma}}{1-\sigma} - \Psi\frac{\left(n_t^s\right)^{1+\gamma}}{1+\gamma}, \qquad (1)$$

where $\sigma, \Psi, \gamma > 0.0$ and χ_t is the AR(1) preference shock, $\chi_t = \rho_\chi \chi_{t-1} + \varepsilon_t^\chi$ with $\varepsilon_t^\chi \sim N(0, \sigma_{\varepsilon^\chi})$, that affects utility from consumption. The bundles of consumption and labor in (1) are obtained using Dixit-Stiglitz aggregators over differentiated consumption goods and labor services, $c_t = \left[\int_0^1 c_t(i)^{\frac{\theta_p - 1}{\theta_p}} di\right]^{\frac{\theta_p}{\theta_p - 1}}$ and $n_t^s = \left[\int_0^1 n_t^s(i)^{\frac{1+\theta_w}{\theta_w}} di\right]^{\frac{\theta_w}{1+\theta_w}}$. The budget constraint for this representative household can be written in nominal terms as follows:

$$\int_0^1 W_t(i) n_t^s(i) di = \int_0^1 P_t(i) c_t(i) di + (1 + R_t)^{-1} B_{t+1} - B_t, \qquad (2)$$

which indicates that labor income on the left-hand side of (2) is spent on purchases of consumption goods and on increasing the amount of risk-free bonds, $(1 + R_t)^{-1} B_{t+1} - B_t$, that yield a nominal interest rate, R_t. Using Dixit-Stiglitz aggregators of the price level and the nominal wage, $P_t = \left[\int_0^1 P_t(i)^{1-\theta_p} di\right]^{\frac{1}{1-\theta_p}}$ and $W_t = \left[\int_0^1 W_t(i)^{1+\theta_w} di\right]^{\frac{1}{1+\theta_w}}$, it can be proved that optimal households' substitutions imply that $\int_0^1 W_t(i) n_t^s(i) di = W_t n_t^s$ and $\int_0^1 P_t(i) c_t(i) di = P_t c_t$. Inserting these results in (2) and dividing by

De Walque, Smets, and Wouters (2006) study the implications of firm-specific labor in a model with Taylor contracts on prices set by firms and wages set by households. Finally, Casares (2007b) also assumes firm-specific labor services in a model with flexible prices and Calvo sticky wages set by either households or firms.

[10] In other words, firms are both monopolistic competitors on providing consumption goods and monopsonistic competitors on demanding labor services.

P_t, the budget constraint in real magnitudes—i.e., units are bundles of consumption goods—becomes

$$\frac{W_t}{P_t} n_t^s = c_t + (1 + R_t)^{-1} \frac{B_{t+1}}{P_t} - \frac{B_t}{P_t}. \tag{3}$$

Using (1) in an infinite time horizon with rational expectations, the representative household wants to maximize $E_t \sum_{j=0}^{\infty} \beta^j U(\chi_{t+j}, c_{t+j}, n_{t+j}^s)$ subject to budget constraints (2) or (3) in period t and future periods.[11] The first-order conditions for the optimal decision on the number of bundles of consumption goods, c_t, and labor services, n_t^s, respectively, are

$$\exp(\chi_t) c_t^{-\sigma} - \xi_t = 0, \tag{4a}$$

$$-\Psi \left(n_t^s\right)^{\gamma} + \xi_t \frac{W_t}{P_t} = 0, \tag{4b}$$

where ξ_t is the Lagrange multiplier of the budget constraint. The value of ξ_t implied by (4a) can be substituted in (4b) and terms can be rearranged to obtain the following labor-supply function for bundles of labor services:

$$n_t^s = \left(\frac{\exp(\chi_t) W_t / P_t}{\Psi c_t^{\sigma}} \right)^{1/\gamma}. \tag{5}$$

Meanwhile, the first-order conditions on the i-th specific type of consumption good and on the i-th specific type of labor service, respectively, are

$$\exp(\chi_t) c_t^{-\sigma} \left(\frac{c_t}{c_t(i)} \right)^{\frac{1}{\theta_p}} - \xi_t \left(\frac{P_t(i)}{P_t} \right) = 0, \tag{6a}$$

$$-\Psi \left(n_t^s\right)^{\gamma} \left(\frac{n_t^s(i)}{n_t^s} \right)^{\frac{1}{\theta_w}} + \xi_t \frac{W_t(i)}{W_t} \frac{W_t}{P_t} = 0. \tag{6b}$$

Combining (4a) and (6a) to eliminate ξ_t leads to the Dixit-Stiglitz demand function

$$c_t(i) = \left(\frac{P_t(i)}{P_t} \right)^{-\theta_p} c_t, \tag{7}$$

[11] As usual, the discount factor is constant at $\beta < 1.0$ and future values are foreseen by applying the rational-expectations operator E_t.

where θ_p is the constant elasticity of substitution for consumption goods. Analogously, a supply curve for the specific type of labor service can be derived by combining (4b) and (6b):

$$n_t^s(i) = \left(\frac{W_t(i)}{W_t}\right)^{\theta_w} n_t^s, \tag{8}$$

in which θ_w is the constant elasticity of substitution across differentiated labor services. Taking logs on both sides of (8) yields

$$\widehat{n}_t^s(i) = \theta_w\big(\widehat{W}_t(i) - \widehat{W}_t\big) + \widehat{n}_t^s, \tag{9}$$

where variables topped with a hat denote the log of the original variable (e.g., $\widehat{W}_t(i) = \log W_t(i)$). The (log-linear) labor-supply curve (9) indicates how households are willing to provide more specific labor services whenever their relative nominal wage increases or, alternatively, whenever there is a higher supply of bundles of labor.

Shifting to labor demand, let us suppose that all firms have access to a production technology with decreasing marginal productivity of labor and a technology shock. Thus, the production function is written for the i-th firm as follows:

$$y_t(i) = \big(\exp(z_t) n_t^d(i)\big)^{1-\alpha}, \tag{10}$$

with $0.0 < \alpha < 1.0$. The amount of firm-specific output, $y_t(i)$, depends on the firm-specific labor demand, $n_t^d(i)$, and on the exogenous AR(1) technology shock, $z_t = \rho_z z_{t-1} + \varepsilon_t^z$ with $\varepsilon_t^z \sim N(0, \sigma_{\varepsilon^z})$. Firms are monopolistic competitors á la Dixit and Stiglitz (1977), with labor demand determined by the level of output. Therefore, we recall the Dixit-Stiglitz demand equation, (7), then use the market-clearing condition, $c_t(i) = y_t(i)$, and the Dixit-Stiglitz output aggregator, $y_t = \left[\int_0^1 y_t(i)^{\frac{\theta_p - 1}{\theta_p}} di\right]^{\frac{\theta_p}{\theta_p - 1}}$, to obtain

$$y_t(i) = \left(\frac{P_t(i)}{P_t}\right)^{-\theta_p} \left[\int_0^1 y_t(i)^{\frac{\theta_p - 1}{\theta_p}} di\right]^{\frac{\theta_p}{\theta_p - 1}}.$$

Then inserting the production function (10), we find

$$\left(\exp(z_t)n_t^d(i)\right)^{1-\alpha} = \left(\frac{P_t(i)}{P_t}\right)^{-\theta_p} \left[\int_0^1 \left(\exp(z_t)n_t^d(i)\right)^{\frac{(1-\alpha)(\theta_p-1)}{\theta_p}} di\right]^{\frac{\theta_p}{\theta_p-1}}.$$

The last expression can be log-linearized to reach the following labor-demand equation:

$$\widehat{n}_t^d(i) = -\frac{\theta_p}{1-\alpha}\left(\widehat{P}_t(i) - \widehat{P}_t\right) + \widehat{n}_t, \tag{11}$$

where $\widehat{n}_t = \int_0^1 \widehat{n}_t^d(i)di$ is the log of aggregate labor.[12]

Now, we are ready to obtain the subordinate nominal wage. As in Bénassy (1995), the wage contract is set at the value that matches labor supply with labor demand. Before introducing wage stickiness, let us examine the wage-setting behavior with fully flexible wages. If all firms can reset their wage contracts every period, the matching condition $\widehat{n}_t^d(i) = \widehat{n}_t^s(i)$, where $\widehat{n}_t^s(i)$ is given by (9) and $\widehat{n}_t^d(i)$ by (11), leads to this (log of) the nominal wage:

$$\widehat{W}_t(i) = \widehat{W}_t - \frac{\theta_p}{\theta_w(1-\alpha)}\left(\widehat{P}_t(i) - \widehat{P}_t\right). \tag{12}$$

There is a negative relationship between the firm-specific optimal price, $\widehat{P}_t(i)$, and the subordinate nominal wage, $\widehat{W}_t(i)$. Those firms that set prices above the aggregate price level will demand less labor and will reduce the nominal wages offered to households in order to reach a perfect match of labor supply with their decreasing demand for labor. With flexible wages, the labor market is in equilibrium because all the pairs of differentiated labor supply and labor demand are well matched.

However, the presence of nominal rigidities on firm-specific wage setting brings in situations of disequilibrium in the labor market regarding the fraction of wage contracts that are not revised. For simplicity, we extend the sticky-price scheme á la Calvo (1983)

[12] Therefore, labor demand becomes the amount of labor effectively employed as typical from a Keynesian economy.

to the resetting of the subordinate wage contracts.[13] Hence, there is a constant probability, η, that the firm is not able to optimally adjust prices, which also causes the lack of wage adjustment. In those situations, next period's prices and nominal wages are left unchanged, assuming that the steady-state rate of inflation is 0 and price/wage indexations do not proceed. With this price and wage stickiness, the wage-setting procedure of labor matching becomes forward looking in order to take into account expectations on future amounts of labor demand and supply attached to the unrevised price and wage. Assuming that the i-th firm can optimize in period t, the labor-matching wage is set at the value that satisfies

$$E_t^{\eta} \sum_{j=0}^{\infty} \beta^j \eta^j \left[\widehat{n}_{t+j}^d(i) - \widehat{n}_{t+j}^s(i)\right] = 0, \tag{13}$$

where E_t^{η} denotes the rational-expectations operator conditional to not being able to change the price and the wage contract in future periods. Inserting the labor-demand equation (9) for any $t+j$ period in (13), and also the labor-supply curve (11) for any $t+j$ period, results in the following (log of) the labor-matching nominal wage:

$$\widehat{W}_t(i) = -\frac{\theta_p}{\theta_w(1-\alpha)}\widehat{P}_t(i) + (1-\beta\eta)E_t \sum_{j=0}^{\infty} \beta^j \eta^j \left(\widehat{W}_{t+j} + \frac{\theta_p}{\theta_w(1-\alpha)}\widehat{P}_{t+j} - \frac{1}{\theta_w}(\widehat{n}_{t+j}^s - \widehat{n}_{t+j})\right) \tag{14}$$

that collapses to (12) in the absence of nominal Calvo-type rigidities ($\eta = 0.0$). As in Blanchard and Galí (2007), let us define the rate of unemployment as follows:[14]

$$u_t = \widehat{n}_t^s - \widehat{n}_t, \tag{15}$$

which can be noticed in (14) referring to the $t+j$ period. As mentioned above, firm-specific sticky wages explain the separation

[13] In the absence of optimal pricing, the firm would have to demand as many labor units as required by the Dixit-Stiglitz demand curve at the current (non-optimal) price, whereas households would have to work that number of hours with no wage revision. Neither party would optimize in their choices of labor.

[14] The rate of unemployment could also have been introduced as $u_t = 1 - \frac{n_t}{n_t^s}$. If so, (15) would be reached by taking logs on both sides of the equivalent expression $1 - u_t = \frac{n_t}{n_t^s}$ and then assuming that $\log(1-u_t) \simeq -u_t$ because u_t is a sufficiently small number.

between the supply of labor bundles and their effective labor demand coming from the unrevised wage contracts. Such an endogenous unemployment is not present in EHL (2000), because the household-specific wage-setting behavior leads to a perfect matching for all pairs of differentiated labor demand and labor supply despite having wage stickiness.

Let us continue the analysis to derive the wage-inflation equation with firm-specific sticky wages. Aggregate wage inflation can be defined as $\pi_t^w = \widehat{W}_t - \widehat{W}_{t-1}$ and, in a similar way, aggregate price inflation as $\pi_t^p = \widehat{P}_t - \widehat{P}_{t-1}$. These definitions allow us to write $\widehat{W}_{t+j} = \widehat{W}_t + \sum_{k=1}^{j} \pi_{t+k}^w$ and $\widehat{P}_{t+j} = \widehat{P}_t + \sum_{k=1}^{j} \pi_{t+k}^p$, which can be inserted into (14) to yield

$$\begin{aligned} \widehat{W}_t(i) - \widehat{W}_t = &-\frac{\theta_p}{\theta_w(1-\alpha)}\big(\widehat{P}_t(i) - \widehat{P}_t\big) - \frac{1-\beta\eta}{\theta_w} E_t \sum_{j=0}^{\infty} \beta^j \eta^j u_{t+j} \\ &+ E_t \sum_{j=1}^{\infty} \beta^j \eta^j \left(\pi_{t+j}^w + \frac{\theta_p}{\theta_w(1-\alpha)} \pi_{t+j}^p \right). \end{aligned} \tag{16}$$

As a well-known result obtained for the Calvo pricing scheme, aggregate price inflation is linked to the log-difference between the optimal price and the aggregate price level:[15]

$$\pi_t^p = \frac{1-\eta}{\eta}\big(\widehat{P}_t(i) - \widehat{P}_t\big). \tag{17}$$

With firm-specific wages, the Calvo fixed-probability scheme for sticky prices also determines the allocation of differentiated wages coming from the (subordinate) sticky wages. Thus, wage inflation is analogously related to the log-difference between the subordinate nominal wage and the aggregate nominal wage,

$$\pi_t^w = \frac{1-\eta}{\eta}\big(\widehat{W}_t(i) - \widehat{W}_t\big). \tag{18}$$

[15] One can reach this result by log-linearizing the aggregate-price-level definition with Calvo-style pricing frictions, $P_t = [(1-\eta)P_t(i)^{1-\theta_p} + \eta P_{t-1}^{1-\theta_p}]^{\frac{1}{1-\theta_p}}$, where P_t is the aggregate price level and $P_t(i)$ is the optimal price.

Combining (16), (17), and (18), we can obtain a relationship between wage inflation and price inflation of this kind:[16]

$$\pi_t^w = \beta E_t \pi_{t+1}^w - \frac{\theta_p}{\theta_w(1-\alpha)}\left(\pi_t^p - \beta E_t \pi_{t+1}^p\right) - \frac{(1-\beta\eta)(1-\eta)}{\eta\theta_w} u_t, \tag{19}$$

which determines wage-inflation fluctuations with firm-specific sticky wages. Using (19) to compute π_{t+j}^w and then substituting the result in (16) yields

$$\widehat{W}_t(i) - \widehat{W}_t = -\frac{\theta_p}{\theta_w(1-\alpha)}\left(\widehat{P}_t(i) - \widehat{P}_t\right) - \frac{1-\beta\eta}{\theta_w} E_t \sum_{j=0}^{\infty} \beta^j u_{t+j}. \tag{20}$$

The relative firm-specific wage contract, $\widehat{W}_t(i) - \widehat{W}_t$, depends on the current relative firm-specific price, $\widehat{P}_t(i) - \widehat{P}_t$, with a negative sign of dependence. If the i-th firm set an optimal price higher than the aggregate price level, its labor demand would fall in response to the decay in the amount of production given by the Dixit-Stiglitz demand curve. Thus, the firm would set a lower nominal-wage contract to clear the supply of labor with its decreasing labor demand. Another determinant of the relative firm-specific wage contract is the expected discounted sum of the rates of unemployment from the current period onward. The influence of unemployment on the labor-matching nominal wage is of a negative sign. A positive unemployment rate means that households wish to work more bundles of labor than the actual number of bundles, which compels a lower nominal wage to satisfy the labor matching condition.

[16] See appendix 1 for the proof. One alternative way to express (19) is

$$\pi_t^w = -\frac{\theta_p}{\theta_w(1-\alpha)}\pi_t^p - \frac{(1-\beta\eta)(1-\eta)}{\eta\theta_w} E_t \sum_{j=0}^{\infty} \beta^j u_{t+j},$$

which means that wage-inflation fluctuations are negatively related to current price inflation and also negatively influenced by the stream of current and expected future rates of unemployment.

3. Pricing and Inflation Dynamics with Firm-Specific Sticky Wages

What are the implications of firm-specific sticky wages on price setting? And what are the implications on price inflation as a result of putting together optimal prices with unchanged prices? This section investigates these questions and derives the Phillips curve for the variant of the New Keynesian model with firm-specific sticky wages (FSW model henceforth). As mentioned above, sticky prices, and the subordinate sticky wages, are jointly introduced by assuming a constant probability for optimal pricing as in Calvo (1983). Therefore, the firm-specific (subordinate) nominal wage (20) is taken into account to find the optimal price because both prices and wages can be simultaneously reset. Supposing that the i-th firm can price optimally in period t, the value of $P_t(i)$ is the one that maximizes conditional intertemporal profits:[17]

$$E_t^{\eta}\sum_{j=0}^{\infty}\Delta_{t,t+j}\eta^j\left[\left(\frac{P_t(i)}{P_{t+j}}\right)^{1-\theta_p}y_{t+j}-\frac{W_t(i)}{P_{t+j}}n_{t+j}^d(i)\right],$$

where the rational-expectations operator, E_t^{η}, is conditional to not being able to change the price and the nominal wage in future periods, and $\Delta_{t,t+j}$ is the stochastic discount factor.[18] The optimality condition for $P_t(i)$ yields

$$E_t^{\eta}\sum_{j=0}^{\infty}\Delta_{t,t+j}\eta^j\left[(1-\theta_P)(P_t(i))^{-\theta_p}(P_{t+j})^{\theta_P-1}y_{t+j}\right.$$
$$\left.-\frac{W_t(i)}{P_{t+j}}\frac{\partial n_{t+j}^d(i)}{\partial y_{t+j}(i)}\frac{\partial y_{t+j}(i)}{\partial P_t(i)}\right]=0. \qquad (21)$$

[17]Notice that total income and labor costs both are expressed in aggregate output units. In addition, total income is obtained as the product of the relative price multiplied by the units of output. So, total income of period t is $\frac{P_t(i)}{P_t}y_t(i)=\frac{P_t(i)}{P_t}(\frac{P_t(i)}{P_t})^{-\theta_p}y_t(i)=(\frac{P_t(i)}{P_t})^{1-\theta_p}y_t(i)$.

[18]The discount factor consistent with the household's optimizing behavior described above is $\Delta_{t,t+j}=\beta^j\frac{\exp(\chi_{t+j})c_{t+j}^{-\sigma}}{\exp(\chi_t)c_t^{-\sigma}}$, which in steady state is constant at β^j.

Meanwhile, the conditional Dixit-Stiglitz demand constraints in any $t+j$ period,

$$y_{t+j}(i) = \left(\frac{P_t(i)}{P_{t+j}}\right)^{-\theta_p} y_{t+j}, \tag{22}$$

imply that

$$\frac{\partial y_{t+j}(i)}{\partial P_t(i)} = -\theta_p (P_t(i))^{-\theta_p - 1} (P_{t+j})^{\theta_P} y_{t+j}. \tag{23}$$

Inserting (23) into (21) and solving out for the optimal price $P_t(i)$, we obtain

$$P_t(i) = \frac{\theta_P}{\theta_P - 1} \frac{E_t^\eta \sum_{j=0}^\infty \Delta_{t,t+j} \eta^j [\psi_{t+j}(i)(P_{t+j})^{\theta_P} y_{t+j}]}{E_t^\eta \sum_{j=0}^\infty \Delta_{t,t+j} \eta^j [(P_{t+j})^{\theta_P - 1} y_{t+j}]}, \tag{24}$$

where $\psi_{t+j}(i) = \frac{W_t(i)}{P_{t+j}} \frac{\partial n_{t+j}^d(i)}{\partial y_{t+j}(i)}$ is the real marginal cost in period $t+j$ subject to the lack of optimal pricing and wage adjustments from period t through period $t+j$. Log-linearizing (24) yields

$$\widehat{P}_t(i) = (1 - \beta\eta) E_t^\eta \sum_{j=0}^\infty \beta^j \eta^j \big(\widehat{P}_{t+j} + \widehat{\psi}_{t+j}(i)\big). \tag{25}$$

The next task is to find an expression for $\widehat{P}_t(i)$ that only depends on aggregate variables so that we can derive a price-inflation equation. The conditional expectation of the log of the firm-specific real marginal cost that appears in (25) can be decomposed as follows:

$$E_t^\eta \widehat{\psi}_{t+j}(i) = \big(\widehat{W}_t(i) - E_t \widehat{P}_{t+j}\big) - E_t^\eta \widehat{mpl}_{t+j}(i), \tag{26}$$

where $\widehat{mpl}_{t+j}(i)$ is the log of the firm-specific marginal product of labor in $t+j$. It should be noticed that $E_t^\eta \widehat{\psi}_{t+j}(i)$ depends on the log of the nominal wage subordinated to the optimal price set in period t, $\widehat{W}_t(i)$. Recalling the production function (10), and the conditional Dixit-Stiglitz demand constraint (22), we can also express $E_t^\eta \widehat{mpl}_{t+j}(i)$ as a function of the optimal price in period t:

$$E_t^\eta \widehat{mpl}_{t+j}(i) = -\frac{\alpha}{1-\alpha} E_t^\eta \widehat{y}_{t+j}(i) = \frac{\theta_p \alpha}{1-\alpha} \big(\widehat{P}_t(i) - E_t \widehat{P}_{t+j}\big) + E_t \widehat{mpl}_{t+j},$$

which can be substituted in (26) to yield

$$E_t^{\eta}\widehat{\psi}_{t+j}(i) = \left(\widehat{W}_t(i) - E_t\widehat{P}_{t+j}\right) - \frac{\theta_p\alpha}{1-\alpha}\left(\widehat{P}_t(i) - E_t\widehat{P}_{t+j}\right) - E_t\widehat{mpl}_{t+j}.$$

We can do some algebra to rewrite the last expression in the following way:[19]

$$E_t^{\eta}\widehat{\psi}_{t+j}(i) = E_t\widehat{\psi}_{t+j} + \left(\widehat{W}_t(i) - \widehat{W}_t - E_t\sum_{k=1}^{j}\pi_{t+k}^{w}\right) - \frac{\theta_p\alpha}{1-\alpha}\left(\widehat{P}_t(i) - \widehat{P}_t - \sum_{k=1}^{j}\pi_{t+k}^{p}\right), \tag{27}$$

where it should be noticed that $\widehat{\psi}_{t+j}$ denotes the log of the aggregate real marginal cost $\widehat{\psi}_{t+j} = (\widehat{W}_{t+j} - \widehat{P}_{t+j}) - \widehat{mpl}_{t+j}$. The relative nominal wage in period t, $\widehat{W}_t(i) - \widehat{W}_t$, is given by equation (20), derived in the previous section. That result can be used in (27) to reach

$$\widehat{\psi}_{t+j}(i) = E_t\widehat{\psi}_{t+j} - \frac{\theta_p(\alpha+\theta_w^{-1})}{1-\alpha}\left(\widehat{P}_t(i) - \widehat{P}_t\right) - \frac{(1-\beta\eta)}{\theta_w}E_t\sum_{j=0}^{\infty}\beta^j u_{t+j} + E_t\sum_{k=1}^{j}\left(\frac{\theta_p\alpha}{1-\alpha}\pi_{t+k} - \pi_{t+k}^{w}\right),$$

which can be substituted in (25) to obtain a value for $\widehat{P}_t(i) - \widehat{P}_t$ that only depends on aggregate variables:

$$\left(1 + \frac{\theta_p(\alpha+\theta_w^{-1})}{1-\alpha}\right)\left(\widehat{P}_t(i) - \widehat{P}_t\right) = (1-\beta\eta)E_t\sum_{j=0}^{\infty}\beta^j\eta^j\widehat{\psi}_{t+j} - \frac{(1-\beta\eta)}{\theta_w}E_t\sum_{j=0}^{\infty}\beta^j u_{t+j} + E_t\sum_{j=1}^{\infty}\beta^j\eta^j\left(\left(1+\frac{\theta_p\alpha}{1-\alpha}\right)\pi_{t+j}^{p} - \pi_{t+j}^{w}\right). \tag{28}$$

[19]First, both $\widehat{W}_{t+j} - \widehat{W}_{t+j}$ and $\widehat{P}_{t+j} - \widehat{P}_{t+j}$ are inserted on the right-hand side of the equation and, secondly, $\widehat{W}_{t+j} = \widehat{W}_t + \sum_{k=1}^{j}\pi_{t+k}^{w}$ and $\widehat{P}_{t+j} = \widehat{P}_t + \sum_{k=1}^{j}\pi_{t+k}^{p}$ are used from the definitions of wage and price inflation.

The terms π^w_{t+j} can be replaced by those obtained when rewriting the expression that appears in footnote (16) for any $t+j$ period. Such substitutions in (28) lead to the following equation:

$$\left(1+\frac{\theta_p(\alpha+\theta_w^{-1})}{1-\alpha}\right)\left(\widehat{P}_t(i)-\widehat{P}_t\right)=(1-\beta\eta)E_t\sum_{j=0}^{\infty}\beta^j\eta^j\widehat{\psi}_{t+j}$$
$$-\frac{1-\beta\eta}{\theta_w}E_t\sum_{j=0}^{\infty}\beta^j u_{t+j}+E_t\sum_{j=1}^{\infty}\beta^j\eta^j\left(\left(1+\frac{\theta_p(\alpha+\theta_w^{-1})}{1-\alpha}\right)\pi^p_{t+j}\right.$$
$$\left.+\frac{(1-\eta)(1-\beta\eta)}{\eta\theta_w}\sum_{k=0}^{\infty}\beta^k u_{t+j+k}\right),$$

where the terms involving unemployment can be rearranged to obtain

$$\left(1+\frac{\theta_p(\alpha+\theta_w^{-1})}{1-\alpha}\right)\left(\widehat{P}_t(i)-\widehat{P}_t\right)=(1-\beta\eta)E_t\sum_{j=0}^{\infty}\beta^j\eta^j\left(\widehat{\psi}_{t+j}-\frac{1}{\theta_w}u_{t+j}\right)$$
$$+E_t\sum_{j=1}^{\infty}\beta^j\eta^j\left(\left(1+\frac{\theta_p(\alpha+\theta_w^{-1})}{1-\alpha}\right)\pi^p_{t+j}\right). \qquad (29)$$

By combining (29) and (17), we can build a dynamic equation for $\pi^p_t-\beta\eta E_t\pi^p_{t+1}$ that, after simplifying terms, results in this price-inflation equation for the FSW model:[20]

$$\pi^p_t=\beta E_t\pi^p_{t+1}+\frac{(1-\eta)(1-\beta\eta)}{\eta\left(1+\frac{\theta_p(\alpha+\theta_w^{-1})}{1-\alpha}\right)}(\widehat{\psi}_t-\frac{1}{\theta_w}u_t). \qquad (30)$$

The dynamic behavior of price inflation in (30) is purely forward looking and resembles quite closely the so-called New Keynesian Phillips curve with flexible wages (Sbordone 2002; Walsh 2003, chap. 5; Woodford 2003, chap. 3) because price inflation depends on the expected future rate of inflation and on the current aggregate

[20]See appendix 2 for the proof.

real marginal costs. However, the introduction of sticky wages set by demand-constrained firms makes the unemployment rate enter (30) with a negative impact on price inflation. The economic intuition behind this influence relies on the connection between prices and wages with firm-specific wage setting. A positive unemployment rate lowers nominal wages (as discussed in section 2), and the subsequent fall in the firm-specific real marginal cost has a negative influence on the rate of price inflation via optimal price resetting.

4. Two Variants for a Sticky-Wage New Keynesian Model

The basic New Keynesian model is built upon three elements: (i) an optimizing IS curve that describes a negative relationship between output and the real interest rate, (ii) a New Keynesian Phillips curve obtained from the aggregation of slowly adjusted prices set by profit-maximizing firms, and (iii) a monetary policy rule that determines short-run changes in the nominal interest rate (McCallum 2001; Walsh 2003, chap. 5). The presence of wage stickiness requires the inclusion of additional equations on the supply side for wage inflation, real marginal costs, or labor productivity as in the EHL (2000) model. Now we investigate the qualitative implications of the firm-specific wage-setting behavior described above in comparison to the common practice of having household-specific sticky wages.[21] Therefore, the analysis compares the structures of the EHL and FSW models.

Firstly, the common parts of the EHL and FSW models are introduced. Following McCallum and Nelson (1999), the IS curve can be obtained from the household's optimizing program described in section 2. The first-order conditions on consumption and bonds can be combined to reach this consumption Euler equation:

$$\frac{\exp(\chi_t)c_t^{-\sigma}}{\beta E_t\big(\exp(\chi_{t+1})c_{t+1}^{-\sigma}\big)} = \frac{1+R_t}{1+E_t\pi_{t+1}^p},$$

[21] Casares (2007b) conducts a similar comparison in a flexible-price scenario.

where inserting the market-clearing equilibrium conditions, $c_t = y_t$ and $c_{t+1} = y_{t+1}$, log-linearizing the result, and recalling the AR(1) generating process for the preference shock yields

$$\widehat{y}_t = E_t\widehat{y}_{t+1} - \frac{1}{\sigma}(R_t - E_t\pi^p_{t+1} - (1-\rho_\chi)\chi_t). \tag{31}$$

The resulting equation (31) is the expectational IS curve that indicates how output fluctuations are forward looking and depend negatively on the real interest rate, $R_t - E_t\pi^p_{t+1}$, and positively on the consumption preference shock, χ_t.

For now, let us suppose that monetary policy is conducted by a central bank whose policy actions follow a Taylor-type monetary policy rule (Taylor 1993) with an interest-rate-smoothing component,

$$R_t = \mu_{\pi^p}\pi^p_t + \mu_{\widetilde{y}}(\widehat{y}_t - \widehat{\overline{y}}_t) + \mu_R R_{t-1}, \tag{32}$$

with μ_{π^p}, $\mu_{\widetilde{y}}$, and $\mu_R \geq 0$ and being numbers that satisfy the Taylor principle to avoid indeterminacy.[22] The output-gap term that appears in (32) is defined as the log-difference between current output and potential (natural-rate) output, $\widehat{y}_t - \widehat{\overline{y}}_t$. Current output is determined by demand conditions in a way depicted by the IS curve (31). Potential output is the amount that would have been produced in the economy if both prices and wages were fully flexible to adjust optimally every period. Dropping nominal rigidities ($\eta = 0.0$), one can find that potential output fluctuations are (exogenously) determined by the following equation:[23]

$$\left(\frac{\alpha+\gamma}{1-\alpha} + \sigma\right)\widehat{\overline{y}}_t = (1+\gamma)z_t + \chi_t. \tag{33}$$

[22]As a general result, determinacy is guaranteed when $\mu_{\pi^p} + \mu_R > 1$. See Woodford (2003, chap. 4) for more detailed discussions.

[23]If prices fully adjust every period, the FSW model collapses into a flexible-price, flexible-wage economy with no unemployment (a kind of RBC economy with imperfect competition and heterogeneous labor). All the firms would choose the same optimal price and the same subordinate nominal wage. Since all the contracts are reset every period, all the pairs of labor supply and labor demand across differentiated labor services are well matched and equal. In turn, the labor market clears in terms of bundles of labor services and the unemployment rate is zero.

Equations (31), (32), and (33) will be shared by the models with either household-specific (EHL model) or firm-specific (FSW model) sticky wages because the wage-setting behavior does not alter their computation. The distinct components of both models are discussed next, mainly regarding the driving forces for price-inflation and wage-inflation fluctuations.

4.1 Firm-Specific Sticky Wages (FSW Model)

The price- and wage-setting interactions described above served to derive equation (30), which governs inflation dynamics in the FSW model. Price inflation is purely forward looking and depends on both the real marginal cost and the rate of unemployment. For comparative purposes with the EHL model, (30) will be transformed into one equivalent expression. Prior to that, let us define the log of the real wage as $\widehat{w}_t = \widehat{W}_t - \widehat{P}_t$ and use it when log-linearizing equation (5) to determine the log of the supply of labor bundles, $\widehat{n}_t^s$, then insert it into equation (15) to obtain the unemployment rate as follows:

$$u_t = \widehat{n}_t^s - \widehat{n}_t = \frac{1}{\gamma}(\widehat{w}_t - \sigma\widehat{y}_t + \chi_t) - \widehat{n}_t.$$

Meanwhile, we have from the utility function (1) that the log of the labor-consumption marginal rate of substitution (MRS) is $\widehat{mrs}_t = \gamma\widehat{n}_t + \sigma\widehat{c}_t - \chi_t$, where inserting the goods market-clearing condition, $\widehat{c}_t = \widehat{y}_t$, and subtracting the log of the real wage result in the following gap between the MRS and the real wage:

$$\widehat{mrs}_t - \widehat{w}_t = \gamma\widehat{n}_t + \sigma\widehat{y}_t - \chi_t - \widehat{w}_t. \tag{34}$$

Interestingly, the FSW model implies a close relationship between u_t and $\widehat{mrs}_t - \widehat{w}_t$. Observing the last two equations, one can see that

$$\widehat{mrs}_t - \widehat{w}_t = -\gamma u_t, \tag{35}$$

which allows us to express the inflation equation (30) in the following manner:

$$\pi_t^p = \beta E_t \pi_{t+1}^p + \frac{(1-\eta)(1-\beta\eta)}{\eta\left(1+\frac{\theta_p(\alpha+\theta_w^{-1})}{1-\alpha}\right)}\left(\frac{1}{\gamma\theta_w}(\widehat{mrs}_t - \widehat{w}_t) - (\widehat{mpl}_t - \widehat{w}_t)\right), \tag{36a}$$

where we also used $\widehat{\psi}_t = \widehat{w}_t - \widehat{mpl}_t$. In turn, the price-inflation dynamics of the FSW model embedded in (36a) are governed by the reaction to two gaps:[24]

(i) the MRS gap, $\widehat{mrs}_t - \widehat{w}_t$, the log-deviation between the marginal rate of substitution and the real wage

(ii) the productivity gap, $\widehat{mpl}_t - \widehat{w}_t$, the log-deviation between the marginal productivity of labor and the real wage

The MRS gap represents the increase in the log of the real wage required to equate the supply of labor bundles to their actual level of employment. The reaction of price inflation to the labor-supply wedge, $\widehat{mrs}_t - \widehat{w}_t$, is absent in the traditional New Keynesian Phillips curve but present in (36a). Thus, the FSW model brings in the MRS gap as another explanatory variable for inflation dynamics due to the combination of sticky prices with firm-specific sticky wages. When the marginal rate of substitution exceeds the real wage, the unemployment rate turns negative, $u_t < 0.0$, because households' supply of labor bundles falls below their actual amount of work. Such negative unemployment has a positive impact on the firm-specific nominal wage subordinated to the optimal price (as discussed in section 2). More costly wages increase real marginal costs, optimal prices are posted higher, and, after aggregation, price inflation will rise.

The productivity gap, $\widehat{mpl}_t - \widehat{w}_t$, enters the inflation equation (36a) with a negative sign to reflect the impact of real marginal costs (note that $\widehat{mpl}_t - \widehat{w}_t$ is the log of the real marginal cost with a minus sign in front). Thus, if labor productivity exceeds the real

[24] Here we are mimicking the interpretation that Walsh (2003, chap. 5) makes from the EHL model.

wage, the real marginal cost becomes negative, and the fraction of firms that are able to reset their prices will post a lower price. So, price inflation falls after a positive productivity gap.

Turning to wage-inflation dynamics in the FSW model, we can substitute the term $\pi_t^p - \beta E_t \pi_{t+1}^p$ implied by (36a) in the wage-inflation equation (19) to yield

$$\pi_t^w = \beta E_t \pi_{t+1}^w + \frac{(1-\eta)(1-\beta\eta)}{\eta\left(1+\alpha\theta_w + \frac{\theta_w(1-\alpha)}{\theta_p}\right)} \left(\frac{1+\alpha(\theta_p - 1)}{\gamma\theta_p} (\widehat{mrs}_t - \widehat{w}_t) + (\widehat{mpl}_t - \widehat{w}_t) \right), \tag{37a}$$

where the unemployment rate was also replaced by its relationship to the MRS gap using equation (35). Both the productivity gap and the MRS gap also affect the rate of wage inflation in the FSW model with a positive influence (the productivity gap had a negative impact on price inflation). The firm-specific wage-setting procedure subordinated to the pricing behavior explains these relationships. Thus, a positive productivity gap, $\widehat{mpl}_t - \widehat{w}_t > 0.0$, implies a negative value for the log of real marginal costs, which would make firms lower optimal prices, increase their amount of output (via a Dixit-Stiglitz demand curve), and thus also increase their labor demand. The (subordinate) firm-specific nominal wage would be raised as necessary to match the increasing labor demand with labor supply. Higher nominal wages on the revised contracts would increase the rate of wage inflation. Concerning the MRS gap, when $\widehat{mrs}_t - \widehat{w}_t > 0$, households wish to work fewer bundles of labor than their current employment, and newly revised nominal-wage contracts will have to be of higher value to match labor supply and labor demand. The fraction of firms that can reset wages would post higher nominal values that on the aggregate would push upward the rate of wage inflation.

Three more equations are needed to close the FSW model (which will also be part of the EHL model). The production function (10) implies this log of the aggregate marginal product of labor:

$$\widehat{mpl}_t = \widehat{y}_t - \widehat{n}_t, \tag{38}$$

and this log of aggregate output:

$$\widehat{y}_t = (1-\alpha)(\widehat{n}_t + z_t), \tag{39}$$

where it should be noticed that labor demand becomes effective labor.[25] Concerning the real-wage dynamics, we can take the first difference on the definition of the log of the real wage, $\widehat{w}_t = \widehat{W}_t - \widehat{P}_t$, to obtain

$$\widehat{w}_t = \widehat{w}_{t-1} + \pi_t^w - \pi_t^p. \tag{40}$$

All in all, the FSW model comprises ten equations, (31)–(40), that may determine solution paths for the ten endogenous variables: π_t^p, π_t^w, $\widehat{w}_t$, $\widehat{mpl}_t$, $\widehat{u}_t$, $\widehat{mrs}_t$, $\widehat{y}_t$, $\widehat{\overline{y}}_t$, $\widehat{n}_t$, and R_t. The model has two predetermined variables ($\widehat{w}_{t-1}$ and R_{t-1}) and two exogenous variables (supply shocks shaping technology, z_t, and demand shocks shaping consumption preference, χ_t).

4.2 Household-Specific Sticky Wages (EHL Model)

Unlike the setup just described, the common practice for a sticky-wage specification in the New Keynesian framework is to let households decide on the nominal-wage contract as first assumed by EHL (2000).[26] They build a labor-market structure with heterogeneous types of labor services, each of them supplied by one differentiated household. Thus, there are household-specific nominal wages that are slowly adjusted with constant probability à la Calvo (1983). Households may be able to set the nominal wage, whereas the amount of labor supplied is labor-demand constraint. Firms employ bundles of labor obtained using a Dixit-Stiglitz aggregator that combines all types of labor services. Thus, firms can substitute between differentiated labor services with a constant elasticity. In turn, EHL (2000) derives the following forward-looking wage-inflation equation:

$$\pi_t^w = \beta E_t \pi_{t+1}^w + \frac{(1-\eta_w)(1-\beta\eta_w)}{\eta_w(1+\gamma\widetilde{\theta}_w)}(\widehat{mrs}_t - \widehat{w}_t). \tag{37b}$$

[25] When firm-specific price and wage contracts are not reoptimized, firms are bound to produce as many units of differentiated output as demanded, whereas households are bound to supply as many units of differentiated labor as required to produce that output demand. Labor demand determines the effective level of employment.

[26] Other recent papers with household-specific sticky wages are Amato and Laubach (2003), Smets and Wouters (2003), Christiano, Eichenbaum, and Evans (2005), and Casares (2007a).

The slope coefficient in (37b) depends on structural parameters such as the Calvo (1983) sticky-wage constant probability, η_w; the elasticity of the labor marginal disutility, γ; the firms' elasticity of substitution across differentiated labor services, $\widetilde{\theta}_w$; and the intertemporal discount parameter, β. The MRS gap is the only driving force on wage-inflation dynamics in the EHL model. The assumption of household-specific sticky wages leaves the wage-inflation fluctuations determined exclusively by variables related to the household sector. Put differently, no firm-related variable such as labor productivity enters the wage-inflation equation in the EHL model.

Similarly for the price-inflation equation, the EHL model does not include the MRS gap, $\widehat{mrs}_t - \widehat{w}_t$, because the price-setting and wage-setting procedures are separated (prices for the firms, wages for the households). When firms set prices, they just look at their real marginal costs and take the same economy-wide nominal wage as given in their pricing decision. In turn, the price-inflation equation of the EHL model can be written as follows:[27]

$$\pi_t^p = \beta E_t \pi_{t+1}^p - \frac{(1-\eta_P)(1-\beta\eta_P)}{\eta_p\left(1+\frac{\alpha\theta_p}{1-\alpha}\right)}(\widehat{mpl}_t - \widehat{w}_t). \tag{36b}$$

To summarize, table 1 reports the determinants of the dynamic behavior of both price and wage inflation in the FSW and EHL models. Productivity gaps, $\widehat{mpl}_t - \widehat{w}_t$, reduce price inflation in both sticky-wage setups, whereas they raise wage inflation only in the FSW model. On the other hand, MRS gaps, $\widehat{mrs}_t - \widehat{w}_t$, have a positive impact on wage inflation in both models and also a negative influence on price inflation in the FSW model.

The set of equations of the EHL model can be obtained by making three changes in the system (31)–(40) belonging to the FSW model. We first introduce (36b) instead of (36a) for the price-inflation dynamics. Secondly, (37b) enters the system, replacing

[27] See Sbordone (2002) for an explicit derivation. The price-inflation equation of the EHL (2000) paper has a slightly different slope coefficient because the real marginal cost is not firm specific in their model.

Table 1. Determinants of Price Inflation and Wage Inflation

Price-Inflation Equation		
	FSW Model, Eq. (36a)	**EHL Model, Eq. (36b)**
Productivity Gap, $\widehat{mpl}_t - \widehat{w}_t$	$(-)$	$(-)$
MRS Gap, $\widehat{mrs}_t - \widehat{w}_t$	$(+)$	0
Wage-Inflation Equation		
	FSW Model, Eq. (37a)	**EHL Model, Eq. (37b)**
Productivity Gap, $\widehat{mpl}_t - \widehat{w}_t$	$(+)$	0
MRS Gap, $\widehat{mrs}_t - \widehat{w}_t$	$(+)$	$(+)$

(37a) for wage-inflation dynamics. The third step would consist of eliminating equation (35) because there is no unemployment in the EHL model. These three variations would lead to a nine-equation system that provides solution paths for the nine endogenous variables: π_t^p, π_t^w, $\widehat{w}_t$, $\widehat{mpl}_t$, $\widehat{mrs}_t$, $\widehat{y}_t$, $\widehat{\overline{y}}_t$, $\widehat{n}_t$, and R_t. Predetermined variables and shocks are assumed to be identical in both sticky-wage setups.

4.3 Baseline Parameterization

In the next two sections, we carry out the business-cycle and monetary policy analysis in the EHL and FSW models. For such applied exercises, some numerical values of their structural parameters are required. Borrowing numbers from the baseline quarterly calibration used in EHL (2000), we set $\beta = 0.99$, $\sigma = 1.5$, $\gamma = 1.5$, $\alpha = 0.3$, $\theta_p = 4.0$, $\widetilde{\theta}_w = 0.4$, and $\eta_p = \eta_w = 0.75$ in the EHL model. These values assigned to β, σ, γ, α, and θ_p are also set in the FSW model. In the FSW model, a single Calvo probability, η, collects the level of price and wage stickiness since wage setting is subordinated to the pricing decision. Thus, in the FSW model, we set the same Calvo probability as in the EHL model, $\eta = 0.75$, which means that both prices and wages are reset optimally once per year.[28] The households' elasticity of substitution regarding the supply of differentiated labor services—exclusive from the FSW model—is set at $\theta_w = 4.0$ also to be equal to the elasticity of substitution of firms' demand for labor in the EHL model.[29] The interest rate monetary policy rule (32) is somehow different here compared with that in EHL (2000), and we assign, on empirical grounds, the Taylor (1993) original coefficients with a significant extent of interest rate smoothing, $\mu_R = 0.8$. Using the partial-adjustment mechanism for monetary policy proposed by Clarida, Galí, and Gertler (1998), the reaction coefficients to inflation deviations and the output gap become $\mu_{\pi^p} = 1.5(1 - 0.8) = 0.3$ and $\mu_{\widetilde{y}} = \frac{0.5}{4}(1 - 0.8) = 0.025$.

As for the stochastic elements, the standard deviations of the innovation of the shocks are chosen with a double criteria. First, total variability of output gives a standard deviation of output equal to 2 percent. Second, supply (technology) shocks account for 60 percent of that output variability in the long-run variance decomposition (100 periods ahead), whereas demand shocks explain the remaining 40 percent.[30] Serial correlation is set to be very high for technology shocks ($\rho_z = 0.95$) and moderately high for demand shocks ($\rho_\chi = 0.80$). Table 2 collects all the baseline numerical values of parameters used in both the FSW and EHL models.

Even though the FSW and EHL models share the same degree of frictions on price and wage setting ($\eta = 0.75$ in the FSW model and $\eta_p = \eta_w = 0.75$ in the EHL model), the slope coefficients in their price-inflation and wage-inflation equations are clearly different (see table 3). Thus, the price-inflation equation in the FSW model has a coefficient on the productivity gap lower than that in the EHL model (0.0207 versus 0.0316). Besides, the

[28] Taylor (1999) reviews a survey of empirical papers to conclude that it can be realistic to assume the same price and wage stickiness in around one optimal adjustment per year.

[29] However, we must keep in mind that these elasticities of substitution have a distinct economic interpretation.

[30] The role of supply and demand shocks in the accounting of output business-cycle fluctuations is a matter of recent controversy. Smets and Wouters (2003, 2007) claim that supply-side shocks originate most output fluctuations in both the United States and the euro area, whereas Dufourt (2005) and Gordon (2005) find that demand shocks explain a fraction of output variability in the United States significantly higher than that due to supply shocks.

Table 2. Baseline Numerical Values of Parameters

FSW Model		EHL Model	
$\beta = 0.99$	$\eta = 0.75$	$\beta = 0.99$	$\eta_p = \eta_w = 0.75$
$\sigma = 1.50$	$\mu_{\pi^p} = 0.30$	$\sigma = 1.50$	$\mu_{\pi^p} = 0.30$
$\gamma = 1.50$	$\mu_{\widetilde{y}} = 0.025$	$\gamma = 1.50$	$\mu_{\widetilde{y}} = 0.025$
$\alpha = 0.30$	$\mu_R = 0.80$	$\alpha = 0.30$	$\mu_R = 0.80$
$\theta_p = 4.00$	$\theta_w = 4.00$	$\theta_p = 4.00$	$\widetilde{\theta}_w = 4.00$
$\rho_z = 0.95$	$\rho_\chi = 0.80$	$\rho_z = 0.95$	$\rho_\chi = 0.80$
$\sigma_{\varepsilon^z} = 1.05\%$	$\sigma_{\varepsilon^\chi} = 1.34\%$	$\sigma_{\varepsilon^z} = 0.98\%$	$\sigma_{\varepsilon^\chi} = 1.32\%$

Table 3. Slope Coefficients at Baseline Values of Parameters

Price-Inflation Equation		
	FSW Model, Eq. (36a)	**EHL Model, Eq. (36b)**
For $\widehat{mpl}_t - \widehat{w}_t$	$\frac{(1-\eta)(1-\beta\eta)}{\eta\left(1+\frac{\theta_p\left(\alpha+\theta_w^{-1}\right)}{1-\alpha}\right)} = 0.0207$	$\frac{(1-\eta_P)(1-\beta\eta_P)}{\eta_p\left(1+\frac{\alpha\theta_p}{1-\alpha}\right)} = 0.0316$
For $\widehat{mrs}_t - \widehat{w}_t$	$\frac{(1-\eta)(1-\beta\eta)}{\eta\left(1+\frac{\theta_p\left(\alpha+\theta_w^{-1}\right)}{1-\alpha}\right)}\frac{1}{\gamma\theta_w} = 0.0035$	0.0
Wage-Inflation Equation		
	FSW Model, Eq. (37a)	**EHL Model, Eq. (37b)**
For $\widehat{mpl}_t - \widehat{w}_t$	$\frac{(1-\eta)(1-\beta\eta)}{\eta\left(1+\alpha\theta_w+\frac{\theta_w(1-\alpha)}{\theta_p}\right)} = 0.0296$	0.0
For $\widehat{mrs}_t - \widehat{w}_t$	$\frac{(1-\eta)(1-\beta\eta)}{\eta\left(1+\alpha\theta_w+\frac{\theta_w(1-\alpha)}{\theta_p}\right)}\frac{1+\alpha(\theta_p-1)}{\gamma\theta_p} = 0.0094$	$\frac{(1-\eta_w)(1-\beta\eta_w)}{\eta_w(1+\gamma\widetilde{\theta}_w)} = 0.0123$

MRS gap has less influence than the productivity gap in price-inflation fluctuations of the FSW model because its coefficient is significantly smaller (0.0035). In the EHL model, there is no effect from the MRS gap on price inflation. Regarding wage-inflation dynamics, the slope coefficients are rather similar for the productivity gap under household-specific or firm-specific sticky wages (0.0123 in the EHL model and 0.0094 in the FSW model). However, wage inflation is more sensitive to the productivity gap than to the MRS gap in the FSW model because its slope coefficient is 0.0296 (more than three times higher). In the EHL model, there is no influence on wage-inflation dynamics coming from the productivity gap.

5. Business-Cycle Analysis

Impulse-response functions can be obtained from innovations in the supply (technology) shock, z_t, and the demand (IS) shock, χ_t.[31] The sizes of the innovations are normalized to one standard deviation (numbers provided in table 2) and are compared in both sticky-wage New Keynesian models. Responses are plotted in figure 1 (supply shock) and figure 2 (demand shock) in percent deviations from the steady-state values for output, the real wage, labor productivity, the marginal rate of substitution, and labor demand, whereas unemployment, price inflation, and wage inflation are directly displayed as level departures from the steady-state rates.

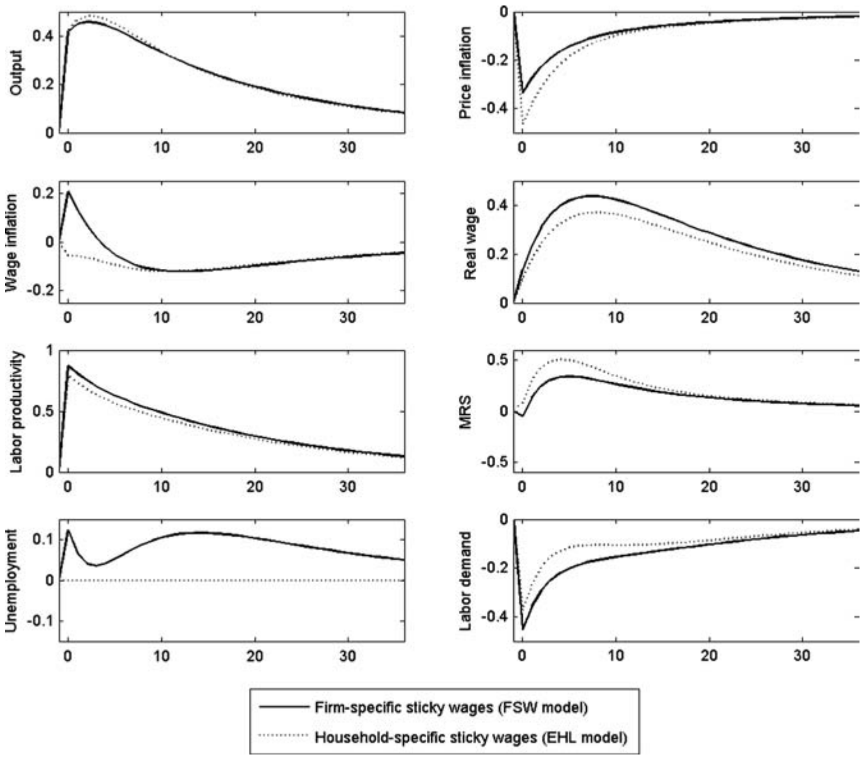
5.1 Supply (Technology) Shock

Figure 1 shows that both output and labor productivity respond to a technology shock with very similar long-lasting rises in the FSW and EHL models. By contrast, wage inflation has a distinctive reaction depending on the wage-setting behavior of the model. When households set nominal wages (EHL model), the model predicts a drop in wage inflation because new wage contracts are set downward. The reason for this behavior is that wage inflation only reacts to the MRS gap (see equation 37b). This gap turns out to be negative due to the initial drop in the MRS and the subsequent increase in the real wage.

If wages are subordinated to the pricing decision of firms (FSW model), wage inflation reacts very differently. In that sticky-wage specification, the change of wage inflation is the result of combining

[31] Even though the demand shock is a consumption-preference shock, we could observe analogous effects from other demand-side shocks such as a fiscal policy shock, an investment-related shock, or a monetary policy shock. Actually, an interest rate shock entering the Taylor-type monetary policy rule (32) would turn absolutely equivalent to a contractionary (negative-signed) demand shock entering the IS curve (31).

Figure 1. One-Standard-Deviation Supply (Technology) Shock: Impulse-Response Functions in the New Keynesian Model with Alternative Sticky-Wage Specifications



the influence of both the productivity gap and the MRS gap. With the technological improvement that brings the shock, the productivity gap becomes clearly positive and outweights the influence of the MRS gap. In turn, wage inflation rises.

The rate of price inflation and the real wage react to the technology shock moving in the same direction in both sticky-wage setups. However, the real wage increases more strongly in the FSW model, whereas price inflation has a more significant drop in the EHL model. The response of the real wage is higher in the FSW model because wage inflation rises there, while it falls in the EHL model. As for the price-inflation reaction, the slope coefficient for the productivity gap is lower in the FSW model (see table 3), as a consequence of the price/wage connections with firm-specific sticky wages. Subsequently, the inflation drop is greater in the EHL model.

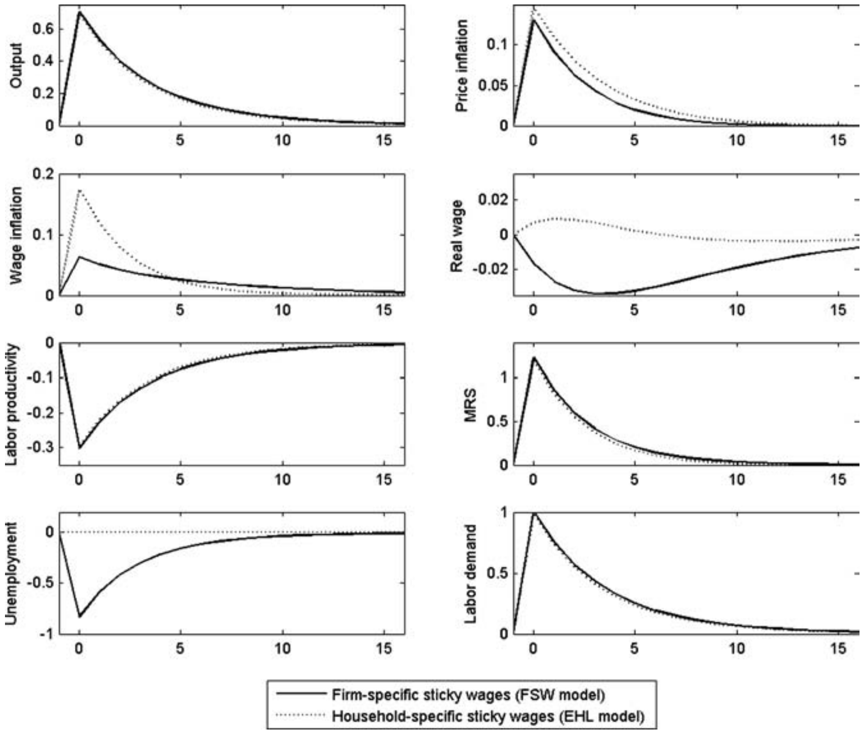
Finally, the rate of unemployment rises in the FSW model, whereas it remains at zero by construction in the EHL model. The unemployment reaction observed in the FSW model is not quantitatively large (the peak increase is approximately one-fourth of the output change). The decline in labor demand that results from the productivity hike explains why unemployment rises. Concretely, the types of labor services that do not have their wage contract revised suffer from a mismatch between labor supply and demand. Their labor demand falls below labor supply because their relative prices are rising due to the lack of price adjustments. Thus, wage stickiness in the FSW model predicts a higher rate of unemployment in response to an expansionary technology shock.

5.2 Demand (IS) Shock

A consumption-preference shock leads to an increase in the households' demand for consumption bundles that expands the IS curve (31) to higher levels of demand-driven output. Figure 2 shows the effects of this expansionary demand shock. At first, output and price inflation rise while (labor) productivity falls under both sticky-wage specifications. The cases of firm-specific or household-specific sticky wages are distinguishable in the reactions of wage inflation, the real wage, and unemployment. Even though wage inflation rises in both sticky-wage setups, the FSW model reports a substantially smaller increase, approximately one-third of that reported in the EHL model (see figure 2). This difference is obtained because in the FSW model wage inflation reacts to the productivity gap, $\widehat{mpl}_t - \widehat{w}_t$, which happens to be negative due to the fall in productivity. In turn, the impact of a positive MRS gap, $\widehat{mrs}_t - \widehat{w}_t > 0$, is partially compensated by a decreasing productivity gap, $\widehat{mpl}_t - \widehat{w}_t < 0$. In the EHL model, there is no influence of productivity on wages, which brings about a stronger reaction of wage inflation.

Meanwhile, price inflation rises in both sticky-wage models due to lower productivity and higher real marginal costs. Quantitatively, the response of price inflation is just slightly higher in the EHL

Figure 2. One-Standard-Deviation Demand (IS) Shock: Impulse-Response Functions in the New Keynesian Model with Alternative Sticky-Wage Specifications



model.[32] As for the real wage, its reaction is obtained when making the difference between the responses of wage inflation and price inflation. In the FSW model, the real wage drops because wage inflation increases to a smaller extent than price inflation. On the contrary, the EHL model reports an increase of wage inflation sufficiently large to produce a higher real wage despite the rise of price inflation.

[32]The responses of price inflation turn out to be nearly identical despite the differences in the driving forces of inflation between the FSW and the EHL model. Thus, the lower slope coefficient in reaction to the productivity gap in the FSW model is almost neutralized by the inflationary effect of the MRS gap that the EHL model does not capture (see table 3 for the numerical values of the slope coefficients).

Hence, the FSW model implies that the real wage would be countercyclical in the presence of a demand shock (higher output, lower real wage), while it reacts in a procyclical fashion in the EHL model (higher output, higher real wage).

Figure 2 also shows that the rate of unemployment falls below its steady-state value in the model with firm-specific sticky wages. Firms demand more labor to produce the additional units of output resulting from the consumption expansion. Since 75 percent of the wage contracts cannot be revised, households have to work longer than desired on those labor services whose contracts are not adjusted. The excess of labor demand over labor supply in aggregate terms represents the reduction of the unemployment rate below its steady-state value.

Summarizing, the impulse-response analysis of the New Keynesian model under different wage-setting behavior confirms a distinctive behavior of the supply side of the model (wage inflation, price inflation, the real wage, and unemployment) in the presence of supply and demand shocks. Wage inflation responds to fluctuations in the MRS if households set nominal wages (EHL model), whereas it reacts to those and also to changes in labor productivity if firms are wage setters (FSW model). In turn, the wage-inflation response to a technology shock is of a different sign. The real wage is procyclical after both shocks in the EHL model, whereas it is procyclical after a supply shock and countercyclical after a demand shock in the FSW model. Regarding inflation, we have observed that the responses are slightly smaller with firm-specific wages. Finally, unemployment rises with a supply shock and falls with a demand shock in the FSW model and has no reaction in the EHL model.

5.3 The Real-Wage Business Cycle and the Sumner-Silver Hypothesis

Sumner and Silver (1989) suggest with an empirical paper that the cyclicality of real wages in the United States depends on the cause of the cycle: if the business cycle is driven by supply shocks, the real wage is strongly procyclical, whereas if demand shocks originate output fluctuations, the real wage becomes clearly anticyclical. Their result provides a convincing empirical explanation of why the correlation between business-cycle fluctuations of output and the real wage is positive and weak in the U.S. economy (Abraham and Haltiwanger 1995). The sign of the correlation varies with the sample period, as the current business cycle is caused by either supply shocks (positive correlation) or demand shocks (negative correlation).

As discussed above, the real wage is procyclical after a technology shock and responds anticyclically in reaction to demand shocks in the FSW model, which replicates the Sumner-Silver empirical findings.[33] By contrast, the reactions of the real wage in the EHL model are procyclical to both supply and demand shocks, which implies that the Sumner-Silver hypothesis cannot be validated and the real wage would turn strongly procyclical.[34]

Table 4 reports the coefficients of correlation between the real wage and output obtained at the baseline price/wage stickiness, $\eta = 0.75$ in the FSW model and $\eta_p = \eta_w = 0.75$ in the EHL model. The real-wage correlations with output are computed in reactions observed to exclusively supply or demand shocks. If households act as wage setters (EHL model), the real wage is highly procyclical with supply shocks, with a coefficient of linear correlation $\rho(\widehat{w}_t, \widehat{y}_t) = 0.92$, and it also shows a clear procyclical behavior with demand shocks, $\rho(\widehat{w}_t, \widehat{y}_t) = 0.71$. The latter fails to be consistent with the Sumner-Silver empirical hypothesis. However, the FSW

Table 4. Real-Wage Correlation with Output at Baseline Price/Wage Stickiness

	FSW Model with $\boldsymbol{\eta = 0.75}$	**EHL Model with $\boldsymbol{\eta_p = \eta_w = 0.75}$**
Supply Shocks	0.94	0.92
Demand Shocks	−0.72	0.71

[33] Bénassy (1995) shows that the cyclicality of the real wage in an optimizing model with flexible prices and predetermined wages is also consistent with the Sumner-Silver hypothesis.

[34] Of course, this result might change if we had other sources of variability (shocks) in the model. Nevertheless, monetary (interest rate) shocks have the same impact as a contractionary demand shock.

model provides real-wage correlations consistent with the Sumner-Silver empirical hypothesis. When supply shocks hit the economy, the real wage is strongly procyclical, $\rho(\widehat{w}_t, \widehat{y}_t) = 0.94$, while the real wage turns clearly anticyclical if demand shocks drive the business cycle, $\rho(\widehat{w}_t, \widehat{y}_t) = -0.72$.

5.3.1 Sensitivity Analysis

Next, we will examine the robustness of the real-wage cyclicality to changes in the level of nominal rigidities on price/wage setting. The analysis is somehow different for each sticky-wage specification at hand. The FSW model has a single Calvo probability for both price and wage stickiness (η). The value of η will be adjusted to consider cases in which the average length of price/wage contracts runs from two quarters ($\eta = 0.5$) to ten quarters ($\eta = 0.9$). The EHL model with household-specific sticky wages features two Calvo probabilities—one for price stickiness affecting firms (η_p) and another one for wage stickiness affecting households (η_w). The sensitivity analysis consists then on moving either η_p or η_w from 0.5 to 0.9 while leaving the other unchanged at 0.75.

Figure 3 displays the results of this sensitivity analysis. The coefficient of correlation between output and the real wage in the presence of supply shocks is always positive and close to 1 at any level of price/wage rigidities in both sticky-wage models (see the plot on the left-hand side of figure 3). Consequently, these correlation coefficients lie on the shaded area that would represent the Sumner-Silver hypothesis.[35] With demand shocks, the real wage is always anticyclical in the FSW model, with negative coefficients of correlation with output that enter the Sumner-Silver area in the plot on the right-hand side of figure 3. Numerical values for the limit cases $\eta = 0.5$ and $\eta = 0.90$ are reported in table 5.

The real-wage cyclicality after a demand shock in the EHL model depends upon the relative price/wage stickiness as shown in figure 3. If the wage-stickiness parameter is at $\eta_w = 0.80$ or higher values, with $\eta_p = 0.75$, wage inflation would barely increase after the shocks and the real-wage response would be more affected by the rise of

[35] Arbitrarily, it is assumed that a coefficient of correlation greater than 0.5 in absolute value represents strong linear dependence.

Figure 3. Real-Wage Cyclicality and Nominal Rigidities: A Robustness Test of the Sumner-Silver Hypothesis in the FSW and EHL Sticky-Wage Models

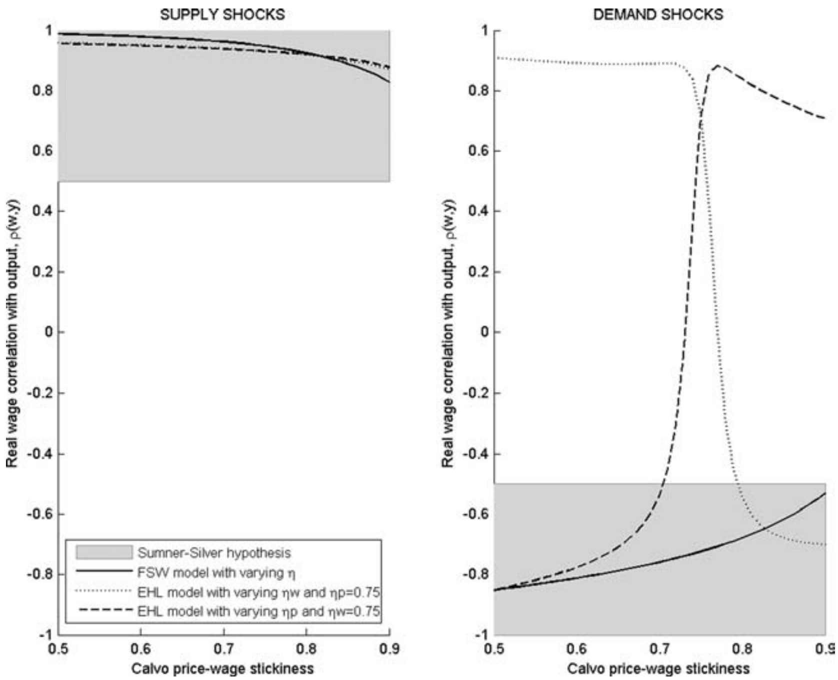


Table 5. Real-Wage Correlation with Output at Various Levels of Price/Wage Stickiness

	FSW Model	EHL Model
Supply Shocks	$\eta = 0.50 \rightarrow 0.98$ $\eta = 0.90 \rightarrow 0.83$	$\eta_p = 0.50$ and $\eta_w = 0.75 \rightarrow 0.95$ $\eta_p = 0.90$ and $\eta_w = 0.75 \rightarrow 0.88$ $\eta_p = 0.75$ and $\eta_w = 0.50 \rightarrow 0.96$ $\eta_p = 0.75$ and $\eta_w = 0.90 \rightarrow 0.87$
Demand Shocks	$\eta = 0.50 \rightarrow -0.85$ $\eta = 0.90 \rightarrow -0.53$	$\eta_p = 0.50$ and $\eta_w = 0.75 \rightarrow -0.85$ $\eta_p = 0.90$ and $\eta_w = 0.75 \rightarrow 0.71$ $\eta_p = 0.75$ and $\eta_w = 0.50 \rightarrow 0.91$ $\eta_p = 0.75$ and $\eta_w = 0.90 \rightarrow -0.70$

price inflation than by that of wage inflation. It results in a real-wage drop and therefore a negative correlation between the real wage and output (see the plot on the right-hand side of figure 3). Table 5 shows that the case when $\eta_p = 0.75$ and $\eta_w = 0.90$ in the EHL model with demand shocks leads to the negative correlation $\rho(\widehat{w}_t, \widehat{y}_t) = -0.70$.

A second possibility is to lower price stickiness. The right-side plot of figure 3 also shows that the coefficient of correlation between the real wage and output with demand shocks turns negative and enters the Sumner-Silver area when η_p falls below 0.70. Thus, table 4 reports significantly countercyclical real wages with demand shocks, $\rho(\widehat{w}_t, \widehat{y}_t) = -0.85$, when setting $\eta_p = 0.50$ and $\eta_w = 0.75$ in the EHL model.

In review, the FSW model satisfies the Sumner-Silver hypothesis at any of the levels of price/wage stickiness examined here. In the EHL model, by contrast, only the calibration with a degree of wage stickiness more persistent than price stickiness is consistent with the Sumner-Silver hypothesis.

6. Monetary Policy Analysis

This section deals with issues related to monetary policy. In particular, we will look for answers to the following two questions:

(i) What are the implications for optimal monetary policy design of having either household-specific or firm-specific sticky wages in the New Keynesian model?

(ii) Can we approximate optimal monetary policy fairly enough using a Taylor-type instrument rule (32) in both sticky-wage cases? If so, what values for the reaction coefficients are required to pursue optimal policy?

So far, monetary policy has followed (32) with a numerical specification for its policy coefficients μ_{π^p}, $\mu_{\widetilde{y}}$, and μ_R that conveys the Taylor (1993) original prescription together with a significant degree of interest rate inertia. Now, we can examine the stabilizing properties of that baseline specification of (32) by comparing it with optimal monetary policy. Furthermore, we will search the *optimized* coefficients for (32) as the triplet that best approximates optimal policy. To begin with, we follow Woodford (2003, chap. 6) and Giannoni and Woodford (2004) to derive the model-based second-order approximation of welfare losses obtained from the utility function of the model. In the FSW model, this welfare-theoretic instantaneous loss function is

$$L_t = \left(\pi_t^p\right)^2 + \lambda(\widetilde{y}_t - \widetilde{y}^*)^2, \tag{41}$$

where $\widetilde{y}_t = \widehat{y}_t - \widehat{\overline{y}}_t$ is the output gap, $\widetilde{y}^* = \frac{1}{\theta_p(\frac{\alpha+\gamma}{1-\alpha}+\sigma)}$ is the steady-state efficient output gap, and the weight on output-gap variability is $\lambda = \frac{(1-\eta)(1-\beta\eta)(\frac{\alpha+\gamma}{1-\alpha}+\sigma)}{\eta(1+\frac{\theta_p(\alpha+\theta_w^{-1})}{1-\alpha})\theta_p}$ (see appendix 3 for its derivation). Therefore, optimal monetary policy targets the variability of price inflation (around its zero steady-state rate) and the variability of the output gap around its steady-state efficient level. This is the same policy recommendation assessed by Woodford (2003, chap. 6) for a New Keynesian model of Calvo-style staggered pricing, heterogeneous labor, and flexible wages. Even though wages are also sticky in the FSW model, the only source of nominal frictions is the Calvo-type probability, η, attached firsthand to the price-setting decision of firms and subsequently to wage adjustments because they are subordinated to optimal prices. In other words, sticky wages and sticky prices are part of the same nominal friction. The role of firm-specific sticky wages in the central-bank loss function is embedded at the value of λ in (41) through the elasticity parameter θ_w, which is absent in models with flexible prices or household-specific sticky wages.[36] Besides, the FSW model permits a trade-off between variabilities of price inflation and the output gap regardless of nominal shocks (documented below), which was not possible in standard models with sticky prices (Taylor 1979; Clarida, Galí, and Gertler 1999).

The welfare-theoretic optimal monetary policy can be obtained in the FSW model by finding the targeting rule that minimizes the expected intertemporal welfare losses, $E_t \sum_{j=0}^{\infty} \beta^j L_{t+j}$, subject to a

[36]In a model with heterogeneous labor, fully flexible wages, and staggered prices á la Calvo, the central bank loss function would be (41) with a slightly different $\lambda = \frac{(1-\eta)(1-\beta\eta)(\frac{\alpha+\gamma}{1-\alpha}+\sigma)}{\eta(1+\frac{\theta_p(\alpha+\gamma)}{1-\alpha})\theta_p}$ (Woodford 2003, chaps. 6 and 8).

reduced set of the structural equations of the model.[37] Such an optimizing program and its first-order conditions are shown in appendix 4. For policy simulations, the baseline numerical parameterization displayed in table 2 can be used to imply the value of $\lambda = 0.021$. This stabilizing policy preference can be used to optimize the policy coefficients for the Taylor-type rule (32). In particular, we search values of the triplet μ_{π^p}, $\mu_{\widetilde{y}}$, and μ_R that minimize the (long-run) unconditional expectation of $\sum_{j=0}^{\infty} \beta^j L_{t+j}$ in the FSW model.[38] It leads to

$$\mu^*_{\pi^p} = 4.09, \ \mu^*_{\widetilde{y}} = 1.17, \text{ and } \mu^*_R = 1.54,$$

which indicate that the nominal interest rate should strongly respond to changes in price inflation and in a more moderate way to the output gap and the previous nominal interest rate. The optimized coefficients are significantly higher than those proposed in Taylor (1993) and used in our previous calibration.

Table 6 examines the stabilizing performance of three alternative monetary policy rules in the FSW model: (i) the (optimal) welfare-theoretic targeting rule with $\lambda = 0.021$; (ii) the baseline Taylor-type rule (32) with $\mu_{\pi^p} = 0.3$, $\mu_{\widetilde{y}} = 0.025$, and $\mu_R = 0.8$; and (iii) the optimized Taylor-type rule (32) with $\mu^*_{\pi^p} = 4.09$, $\mu^*_{\widetilde{y}} = 1.17$, and $\mu^*_R = 1.54$. One can see that the standard deviations of both price inflation and the output gap are much lower when applying the optimal policy compared with the baseline instrument rule. They are cut to approximately one-fourth of their values—from 0.68 percent to 0.16 percent in the case of price inflation, and from 0.67 percent to 0.19 percent in the output gap. However, both wage inflation and the nominal interest rate report a somewhat higher standard deviation with the welfare-theoretic targeting rule because optimal policy does not contemplate any concern on their variabilities (see table 6 for the numbers). When switching from the baseline to the optimized coefficients in (32), the standard deviation of price inflation falls to 0.18 percent and that

[37]The optimal monetary policy analysis based on the utility function is a subproduct of the targeting-rules approach introduced by Svensson (1999) and Woodford (1999).

[38]This same criterion has been used for monetary policy analysis by Levin and Williams (2003), Adalid et al. (2005), and Casares (2007a).

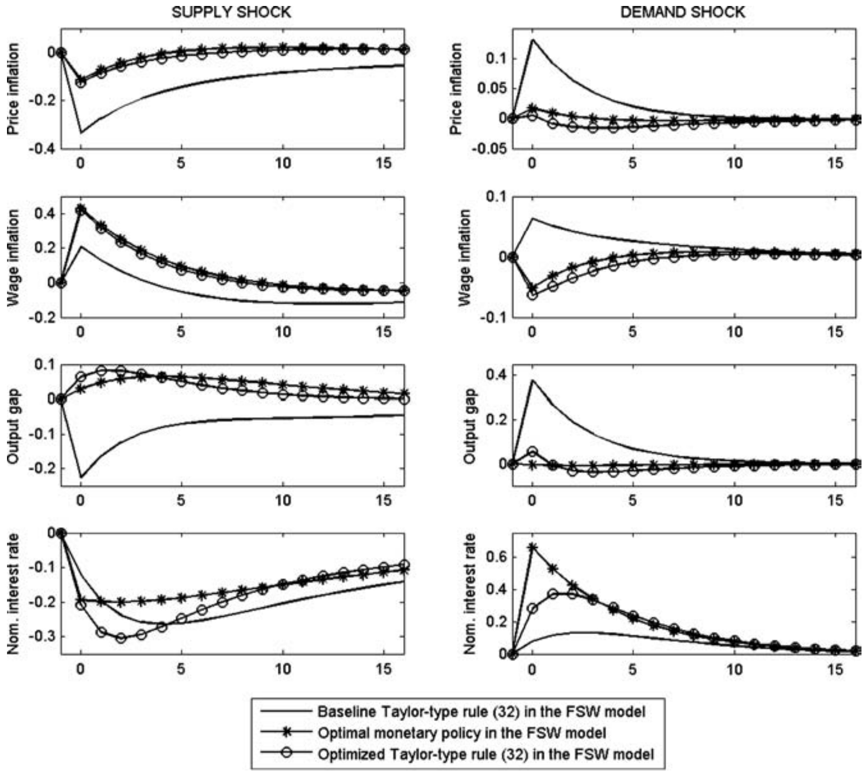
Table 6. Performance of Monetary Policy Rules

	Std. Deviations (%, Annualized)				Loss Ratio
FSW Model	$\boldsymbol{\pi^p}$	$\boldsymbol{\pi^w}$	$\boldsymbol{\widetilde{y}}$	$\boldsymbol{R}$	$\boldsymbol{L/L^*}$
Welfare-Theoretic Targeting Rule (Optimal) ($\lambda = 0.021$)	0.16	0.69	0.19	1.33	1.00
Baseline Taylor-Type Rule (32) ($\mu_{\pi^p} = 0.30$, $\mu_{\widetilde{y}} = 0.025$, $\mu_R = 0.80$)	0.68	0.60	0.67	1.01	16.64
Optimized Taylor-Type Rule (32) ($\mu^*_{\pi^p} = 4.09$, $\mu^*_{\widetilde{y}} = 1.17$, $\mu^*_R = 1.54$)	0.18	0.65	0.20	1.21	1.29
EHL Model	$\boldsymbol{\pi^p}$	$\boldsymbol{\pi^w}$	$\boldsymbol{\widetilde{y}}$	$\boldsymbol{R}$	$\boldsymbol{\mathcal{L}/\mathcal{L}^*}$
Welfare-Theoretic Targeting Rule (Optimal) ($\lambda_{\pi^p} = 0.357$, $\lambda_{\widetilde{y}} = 0.012$)	0.50	0.20	0.03	1.40	1.00
Baseline Taylor-Type Rule (32) ($\mu_{\pi^p} = 0.30$, $\mu_{\widetilde{y}} = 0.025$, $\mu_R = 0.80$)	0.89	0.50	0.57	1.16	4.96
Optimized Taylor-Type Rule (32) ($\mu^*_{\pi^p} = 39.46$, $\mu^*_{\widetilde{y}} = 100.0$, $\mu^*_R = 0.72$)	0.50	0.22	0.05	1.49	1.04

of the output gap to 0.20 percent, which are values only slightly higher than the numbers obtained under the optimal policy. In overall terms, the ratio of the unconditional loss value under the Taylor-type rule (32) divided by that loss under the optimal policy (denoted as L/L^* in table 6) is more than 16 with the baseline coefficients in (32) and gets reduced to only 1.29 with the optimized coefficients in (32). Therefore, it could be said that the stabilizing performance of the instrument rule (32) with the baseline coefficients is quite poor, whereas replacing those with optimized coefficients provides a very good approximation to optimal policy.

Figure 4 shows how the responses of price inflation and the output gap to supply and demand shocks in the FSW model are

Figure 4. Monetary Policy Analysis in the FSW Model: Responses to a Supply Shock (Left) and to a Demand Shock (Right) under Different Monetary Policy Rules



quantitatively much smaller under the optimal welfare-theoretic policy compared with the baseline specification of (32). By contrast, the use of the optimized coefficients in (32) allows that Taylor-type rule to mimic fairly well the responses of price inflation and the output gap obtained with the optimal policy, which confirms its good stabilizing performance. Wage inflation responds more aggressively under the optimal policy (and the optimized Taylor-type rule) because the optimal monetary policy is not aimed at stabilizing wages. Figure 4 also displays the reactions of the nominal interest rate under optimal monetary policy. If there is a supply shock, the nominal interest rate falls in a way similar to that implied by the baseline Taylor-type rule. However, the reaction to a demand shock is much more aggressive with the optimal policy since the nominal interest rate is raised four or five times higher than the level reached with the baseline Taylor-type rule. Such a severe policy tightening leads to a demand contraction that nearly neutralizes the initial expansionary shock. Subsequently, the output gap is practically erased and price inflation stays near 0.

Let us turn to the EHL model for a comparison of optimal monetary policy with household-specific sticky wages. The welfare-theoretic loss function of the EHL model was already obtained by Woodford (2003, chap. 6) and Giannoni and Woodford (2004) as a weighted average of variabilities involving price inflation, wage inflation, and the output gap:

$$\mathcal{L}_t = \lambda_{\pi^p}\left(\pi_t^p\right)^2 + (1-\lambda_{\pi^p})\left(\pi_t^w\right)^2 + \lambda_{\widetilde{y}}(\widetilde{y}_t - \widetilde{y}^*)^2, \tag{42}$$

where the weights on the policy targets are $\lambda_{\pi^p} = \frac{\theta_p \kappa_p^{-1}}{\theta_p \kappa_p^{-1} + \widetilde{\theta}_w (1-\alpha)\kappa_w^{-1}}$ and $\lambda_{\widetilde{y}} = \frac{\left(\frac{\alpha+\gamma}{1-\alpha}+\sigma\right)}{\theta_p \kappa_p^{-1} + \widetilde{\theta}_w (1-\alpha)\kappa_w^{-1}}$, with κ_p and κ_w, respectively, denoting the slope coefficients in the price-inflation and wage-inflation equations, (36b) and (37b). Thus, if households are the wage-setting actors, the optimal monetary policy targets wage-inflation variability, which was not included in the loss function of the FSW model. With the numerical values assigned in the baseline calibration (table 2), the stabilizing policy weights in (42) are $\lambda_{\pi^p} = 0.357$ and $\lambda_{\widetilde{y}} = 0.012$. The price-inflation weight indicates that optimal policy should be more oriented to fighting volatility of wage inflation than of price inflation.[39]

The welfare-theoretic targeting rule for the EHL model was derived in appendix 5 by minimizing $E_t \sum_{j=0}^{\infty} \beta^j \mathcal{L}_{t+j}$ subject to the structural equations of the model. Table 6 shows its stabilizing performance. Despite the apparently low value of $\lambda_{\widetilde{y}}$, optimal policy in the EHL model puts the economy very close to the (natural-rate) frictionless scenario because the output gap barely fluctuates. Thus,

[39]This is in deep contrast to the much higher number for λ_{π^p} suggested by Giannoni and Woodford (2004) for the reasons discussed in Casares (2007a).

the standard deviation of the output gap is only 0.03 percent under the welfare-theoretic optimal policy, which seems very low compared with the value of 0.57 percent obtained under the baseline Taylor-type rule (32). Meanwhile, the variabilities of price inflation and wage inflation get moderate reductions when implementing the optimal policy compared with the baseline Taylor-type rule (32). The standard deviation of price inflation falls from 0.89 percent to 0.50 percent and that of wage inflation decreases from 0.50 percent to 0.20 percent.

Using the optimal welfare-theoretic monetary policy and the same criterion mentioned for the case of the FSW model, the optimized coefficients on the Taylor-type rule (32) of the EHL model are[40]

$$\mu^*_{\pi^p} = 39.46, \ \mu^*_{\tilde{y}} = 100.0, \ \text{and} \ \mu^*_R = 0.72,$$

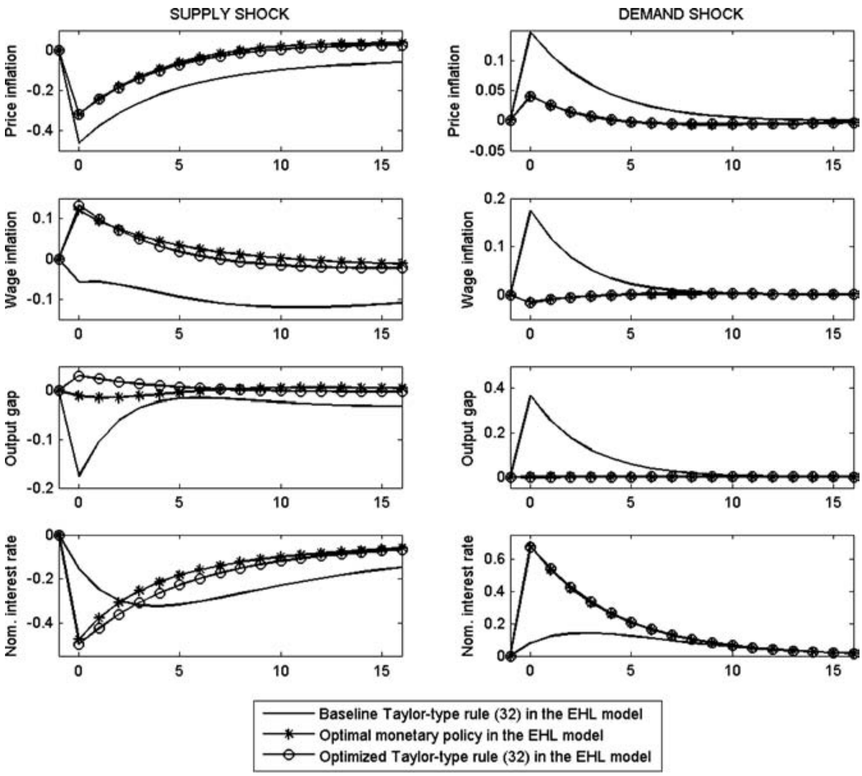
which clearly reflect the major concern of stabilizing the output gap due to its large reaction coefficient, $\mu^*_{\tilde{y}} = 100.0$.

The standard deviations of the targeted variables in the EHL model with the optimized Taylor-type rule (32) are very similar to the numbers obtained with the welfare-theoretic targeting rule. Actually, price inflation has the same volatility, whereas wage inflation and the output gap report only slightly higher numbers (see bottom part of table 6). Unlike the FSW model, the presence of wage inflation in the loss function (42) causes the optimal monetary policy to significantly reduce the wage-inflation volatility. In terms of welfare losses, the ratio of unconditional welfare losses is almost 5 with the baseline calibration of (32), and that can be reduced to only 1.06 (i.e., just a 6 percent higher loss) with the optimized coefficients in (32). Therefore, the Taylor-type rule (32) with $\mu^*_{\pi^p} = 39.46$, $\mu^*_{\tilde{y}} = 100.0$, and $\mu^*_R = 0.72$ approximates fairly well the stabilizing performance achieved with optimal monetary policy.

Impulse-response functions in the EHL model obtained under the three monetary policy rules at hand are shown in figure 5.

[40]The computation of the optimized coefficients was taken with an upper-level bound at 100.0 to avoid excessively large numbers.

Figure 5. Monetary Policy Analysis in the EHL Model: Responses to a Supply Shock (Left) and to a Demand Shock (Right) under Different Monetary Policy Rules



The implementation of the optimized Taylor-type rule (32) leads to responses of the three targeted variables (price inflation, wage inflation, and the output gap) that are very similar to those responses obtained under the optimal monetary policy. The responses under the baseline Taylor-type rule are clearly of larger magnitudes, especially in the case of the output gap. A supply-side technology shock (left side of figure 5) leads to an immediate fall in the nominal interest rate under the optimal monetary policy in order to stimulate output via demand and turn the output gap to the positive side.

As shown on the right side of figure 5, the demand shock also leads to an aggressive interest rate reaction under the welfare-theoretic optimal rule of the EHL model. The expansionary shift on the IS curve is almost neutralized by the combination of a higher nominal interest rate with sticky prices that raises the real interest rate. In turn, the positive output gap that was obtained with the baseline Taylor-type rule is swept away, as also occurred in the FSW model. Meanwhile, price inflation and, especially, wage inflation also show much smaller reactions after the shocks when implementing optimal monetary policy.

Comparing the optimal policies displayed in figures 4 and 5, the responses of wage inflation are significantly smaller in the EHL model than in the FSW model with either demand and supply shocks, which reflects the policy preference for stabilizing wage inflation embedded in (42). On the contrary, price inflation responds more strongly to both shocks in the EHL model, because the optimal policy in the FSW model shows a stronger concern for price-inflation stabilization. The output gap shows a larger reaction after a supply shock in the FSW model, whereas it is practically eliminated after a demand shock in both models.

7. Conclusions

This paper shows that the assumption of who set wages (firms or households) in the New Keynesian model with sticky wages is not trivial. If there are firm-specific sticky wages (FSW model), the wage-setting decision depends on the specific pricing conditions of the firm, whereas with household-specific sticky wages, prices and wages are set independently (as in the EHL model).

Several consequences emerge from the price/wage interactions of firm-specific sticky wages. First, the labor market of the FSW model delivers an endogenous measure of unemployment, which was absent in the prominent EHL model. Only the fraction of wage contracts reset over the current period provide a matching between differentiated labor supply and labor demand. The remaining fraction of non-adjusted wages bring disequilibrium between pairs of labor demand and labor supply that, after aggregation, provide endogenous unemployment. Therefore, introducing sticky wages set by firms serves to incorporate unemployment in a New Keynesian model in an alternative way to the presence of search and matching frictions (e.g., Christoffel and Linzert 2005).

Secondly, firm-specific sticky wages have qualitative implications on both price-inflation and wage-inflation dynamics. The gaps of the real wage with respect to labor productivity and the households' marginal rate of substitution (MRS) are the driving forces for fluctuations on both variables in the FSW model. In the EHL model, by contrast, there is a separation between the productivity gap, which only affects price-inflation fluctuations, and the MRS gap, which only determines wage-inflation fluctuations.

Impulse-response functions indicate that output reacts to supply and demand shocks in similar patterns under both sticky-wage specifications. However, the introduction of firm-specific sticky wages becomes crucial for other macroeconomic variables. In addition to the key issue of the absence or presence of unemployment, variables related to the labor market such as wage inflation or the real wage have shown distinctive business-cycle patterns. For example, the FSW model predicts procyclical reactions of the real wage to supply shocks and countercyclical real-wage responses to demand shocks, which is consistent with the empirical arguments pointed out by Sumner and Silver (1989) to explain the mildly procyclical real wages observed in the United States. The Sumner-Silver hypothesis is replicated in the EHL model only in cases when wage stickiness is more persistent than price stickiness.

Finally, firm-specific or household-specific sticky wages also matter for monetary policy analysis. The (welfare-theoretic) optimal monetary policy in the FSW model targets variabilities of price inflation and the output gap, whereas in the EHL model the optimal policy targets are price inflation, wage inflation, and the output gap. In comparative terms, volatilities of wage inflation and the output gap are higher in the optimal monetary policy of the FSW model. By contrast, price inflation has a higher variability in the EHL model. With both sticky-wage specifications, a Taylor-type rule for the nominal interest rate with the original Taylor (1993) coefficients, along with a significant interest-rate-smoothing component, provides a poor approximation to optimal monetary policy. However, the coefficients of such an instrument rule can be optimized with a welfare-based criterion to reach a good stabilizing performance that closely approximates optimal policy.

Appendix 1. Derivation of the Dynamic Relationship between Wage Inflation and Price Inflation, Equation (19), in the Model with Firm-Specific Sticky Wages

The paper shows in section 2 that the subordinate relative wage with Calvo-style rigidities can be written in log-linear terms as follows:

$$\widehat{W}_t(i) - \widehat{W}_t = -\frac{\theta_p}{\theta_w(1-\alpha)}\left(\widehat{P}_t(i) - \widehat{P}_t\right) - \frac{1-\beta\eta}{\theta_w}E_t\sum_{j=0}^{\infty}\beta^j\eta^j u_{t+j}$$
$$+ E_t\sum_{j=1}^{\infty}\beta^j\eta^j\left(\pi^w_{t+j} + \frac{\theta_p}{\theta_w(1-\alpha)}\pi^p_{t+j}\right),$$

which corresponds to equation (16) of the main text. Substituting the relative price and the relative wage for their respective counterparts in terms of price inflation and wage inflation, $\widehat{P}_t(i) - \widehat{P}_t = \frac{\eta}{1-\eta}\pi^p_t$ and $\widehat{W}_t(i) - \widehat{W}_t = \frac{\eta}{1-\eta}\pi^w_t$, we obtain

$$\pi^w_t = -\frac{\theta_p}{\theta_w(1-\alpha)}\pi^p_t - \frac{(1-\beta\eta)(1-\eta)}{\eta\theta_w}E_t\sum_{j=0}^{\infty}\beta^j\eta^j u_{t+j}$$
$$+ \frac{1-\eta}{\eta}E_t\sum_{j=1}^{\infty}\beta^j\eta^j\left(\pi^w_{t+j} + \frac{\theta_p}{\theta_w(1-\alpha)}\pi^p_{t+j}\right).$$

Moving one period forward from the last expression, we can compute $\beta\eta E_t\pi^w_{t+1}$ and then notice that

$$\pi^w_t - \beta\eta E_t\pi^w_{t+1} = -\frac{\theta_p}{\theta_w(1-\alpha)}\left(\pi^p_t - \beta\eta E_t\pi^p_{t+1}\right) - \frac{(1-\beta\eta)(1-\eta)}{\eta\theta_w}u_t$$
$$+ \frac{1-\eta}{\eta}\beta\eta E_t\left(\pi^w_{t+1} + \frac{\theta_p}{\theta_w(1-\alpha)}\pi^p_{t+1}\right),$$

which simplifies to equation (19) in the main text:

$$\pi^w_t = \beta E_t\pi^w_{t+1} - \frac{\theta_p}{\theta_w(1-\alpha)}\left(\pi^p_t - \beta E_t\pi^p_{t+1}\right) - \frac{(1-\beta\eta)(1-\eta)}{\eta\theta_w}u_t.$$

After recursive substitutions of expected next period's wage inflation, we can also obtain this alternative expression of wage inflation:

$$\pi_t^w = -\frac{\theta_p}{\theta_w(1-\alpha)}\pi_t^p - \frac{(1-\eta)(1-\beta\eta)}{\eta\theta_w}E_t\sum_{j=0}^{\infty}\beta^j u_{t+j}.$$

Appendix 2. From the (Log-Linear) Optimal Price (29) to the Phillips Curve (30) in the Model with Firm-Specific Sticky Wages

The log-difference between the optimal price and the aggregate price level is given by equation (29) of the main text,

$$\left(1+\frac{\theta_p(\alpha+\theta_w^{-1})}{1-\alpha}\right)\left(\widehat{P}_t(i)-\widehat{P}_t\right) = (1-\beta\eta)E_t\sum_{j=0}^{\infty}\beta^j\eta^j\left(\widehat{\psi}_{t+j}-\frac{1}{\theta_w}u_{t+j}\right)$$
$$+E_t\sum_{j=1}^{\infty}\beta^j\eta^j\left(\left(1+\frac{\theta_p(\alpha+\theta_w^{-1})}{1-\alpha}\right)\pi_{t+j}^p\right),$$

where we can insert $\widehat{P}_t(i)-\widehat{P}_t = \frac{\eta}{1-\eta}\pi_t^p$ from (17) to yield

$$\pi_t^p = \frac{(1-\eta)(1-\beta\eta)}{\eta\left(1+\frac{\theta_p(\alpha+\theta_w^{-1})}{1-\alpha}\right)}E_t\sum_{j=0}^{\infty}\beta^j\eta^j\left(\widehat{\psi}_{t+j}-\frac{1}{\theta_w}u_{t+j}\right)$$
$$+\frac{1-\eta}{\eta}E_t\sum_{j=1}^{\infty}\beta^j\eta^j\pi_{t+j}^p.$$

If we move one period forward from price inflation, premultiply the result by $\beta\eta$, and apply the rational-expectations operator in period t, we obtain

$$\beta\eta E_t\pi_{t+1}^p = \frac{(1-\eta)(1-\beta\eta)}{\eta\left(1+\frac{\theta_p(\alpha+\theta_w^{-1})}{1-\alpha}\right)}E_t\sum_{j=1}^{\infty}\beta^j\eta^j\left(\widehat{\psi}_{t+j}-\frac{1}{\theta_w}u_{t+j}\right)$$
$$+\frac{1-\eta}{\eta}E_t\sum_{j=2}^{\infty}\beta^j\eta^j\pi_{t+j}^p.$$

Using the last two expressions, we compute $\pi_t^p - \beta\eta E_t\pi_{t+1}^p$ to find

$$\pi_t^p - \beta\eta E_t\pi_{t+1}^p = \frac{(1-\eta)(1-\beta\eta)}{\eta\left(1+\frac{\theta_p(\alpha+\theta_w^{-1})}{1-\alpha}\right)}\left(\widehat{\psi}_t - \frac{1}{\theta_w}u_t\right) + \frac{1-\eta}{\eta}\beta\eta E_t\pi_{t+1}^p,$$

which reduces to the New Keynesian Phillips curve (30),

$$\pi_t^p = \beta E_t\pi_{t+1}^p + \frac{(1-\eta)(1-\beta\eta)}{\eta\left(1+\frac{\theta_p(\alpha+\theta_w^{-1})}{1-\alpha}\right)}\left(\widehat{\psi}_t - \frac{1}{\theta_w}u_t\right).$$

Appendix 3. Derivation of the Welfare-Theoretic Loss Function in a New Keynesian Model with Firm-Specific Sticky Wages (FSW Model)

All the approximations taken here are based on second-order Taylor expansions used in Erceg, Henderson, and Levin (2000, 307–12) and Woodford (2003, 692–96).

In the New Keynesian model with wage-setting firms, the household utility function and the social utility function are the same because households are alike. Moreover, the effective amount of bundles of labor services is given by demand conditions. Under this assumption, the values of the utility function (1) in period t can be rewritten in effective terms as follows:

$$U_t = \exp(\chi_t)\frac{c_t^{1-\sigma}}{1-\sigma} - \Psi\frac{n_t^{1+\gamma}}{1+\gamma}, \tag{43}$$

where n_t is the number of effective bundles of labor obtained from the aggregation of the demand for labor services over the continuum of firms:

$$n_t = \left[\int_0^1 n_t^d(i)^{\frac{1+\theta_w}{\theta_w}}di\right]^{\frac{\theta_w}{1+\theta_w}}. \tag{44}$$

Let us separate (43) into its two terms: $V_t = \exp(\chi_t)\frac{c_t^{1-\sigma}}{1-\sigma}$ and $S_t = \Psi\frac{(n_t)^{1+\gamma}}{1+\gamma}$. They can be approximated by the Taylor-series expansions

$$V_t \simeq yU_c\left(\widehat{y}_t + \frac{1}{2}(1-\sigma)\widehat{y}_t^2 + \widehat{y}_t\chi_t\right) \text{ and} \tag{45a}$$

$$S_t \simeq nU_n \left(\widehat{n}_t + \frac{1}{2}(1+\gamma)\widehat{n}_t^2 \right), \tag{45b}$$

where the market-clearing condition $\widehat{c}_t = \widehat{y}_t$ was already used in (45a). The term on disutility from labor bundles, S_t, will be affected by the degree of price/wage dispersion through its impact on fluctuations of the demand for labor bundles, $\widehat{n}_t$. Thus, a Taylor-series expansion on (44) yields

$$\widehat{n}_t \simeq E_i \widehat{n}_t^d(i) + \frac{1}{2}\left(\frac{1+\theta_w}{\theta_w}\right) var_i \widehat{n}_t^d(i), \tag{46}$$

where $E_i \widehat{n}_t^d(i)$ and $var_i \widehat{n}_t^d(i)$, respectively, denote the expected value of the demand for labor and its variance computed across the differentiated firms. Recalling the log-linear production function (10) for the specific i-th firm, we obtain

$$\widehat{n}_t^d(i) = (1-\alpha)^{-1}\widehat{y}_t(i) - z_t,$$

which, aggregating over the i space, implies

$$E_i \widehat{n}_t^d(i) = (1-\alpha)^{-1} E_i \widehat{y}_t(i) - z_t, \tag{47a}$$

$$var_i \widehat{n}_t^d(i) = (1-\alpha)^{-2} var_i \widehat{y}_t(i). \tag{47b}$$

The Dixit-Stiglitz demand function (7) implies that in equilibrium

$$y_t(i) = \left(\frac{P_t(i)}{P_t}\right)^{-\theta_p} y_t, \tag{48}$$

which leads to the Taylor-series approximation for output fluctuations:

$$\widehat{y}_t \simeq E_i \widehat{y}_t(i) + \frac{1}{2}\frac{\theta_p - 1}{\theta_p} var_i \widehat{y}_t(i). \tag{49}$$

Substituting the pair (47a) and (47b) into (46), we obtain

$$\widehat{n}_t \simeq (1-\alpha)^{-1} E_i \widehat{y}_t(i) - z_t + \frac{1}{2}\left(\frac{1+\theta_w}{\theta_w}\right)(1-\alpha)^{-2} var_i \widehat{y}_t(i),$$

where inserting $E_i\widehat{y}_t(i) = \widehat{y}_t - \frac{1}{2}\frac{\theta_p - 1}{\theta_p} var_i\widehat{y}_t(i)$ from (49) yields

$$\widehat{n}_t \simeq (1-\alpha)^{-1}\left(\widehat{y}_t - \frac{1}{2}\frac{\theta_p - 1}{\theta_p} var_i\widehat{y}_t(i)\right) \\ - z_t + \frac{1}{2}\left(\frac{1+\theta_w}{\theta_w}\right)(1-\alpha)^{-2} var_i\widehat{y}_t(i).$$

Putting the terms involving $var_i\widehat{y}_t(i)$ together and dropping the exogenous variable z_t, we get

$$\widehat{n}_t \simeq (1-\alpha)^{-1}\widehat{y}_t + \frac{1}{2}(1-\alpha)^{-1}\left(\frac{\alpha + \theta_w^{-1}}{1-\alpha} + \frac{1}{\theta_p}\right) var_i\widehat{y}_t(i). \quad (50)$$

The substitution of (50) into (45b) yields

$$S_t \simeq nU_n\left((1-\alpha)^{-1}\widehat{y}_t + \frac{1}{2}(1-\alpha)^{-1}\left(\frac{\alpha + \theta_w^{-1}}{1-\alpha} + \frac{1}{\theta_p}\right) var_i\widehat{y}_t(i) \right. \\ \left. + \frac{1}{2}(1+\gamma)\widehat{n}_t^2\right),$$

where using $\widehat{n}_t^2 = (1-\alpha)^{-2}\widehat{y}_t^2 - 2(1-\alpha)^{-1}\widehat{y}_t z_t + z_t^2$ from the log-linear production function (10) results in

$$S_t \simeq nU_n\left((1-\alpha)^{-1}\widehat{y}_t + \frac{1}{2}(1-\alpha)^{-1}\left(\frac{\alpha + \theta_w^{-1}}{1-\alpha} + \frac{1}{\theta_p}\right) var_i\widehat{y}_t(i)\right) \\ + nU_n\left(\frac{1}{2}(1+\gamma)\big((1-\alpha)^{-2}\widehat{y}_t^2 - 2(1-\alpha)^{-1}\widehat{y}_t z_t + z_t^2\big)\right). \quad (51)$$

Following Woodford (2003, chap. 6), the steady-state solution of the model implies the relationship $nU_n = yU_c(1-\alpha)(1-\Phi_y)$, where $1-\Phi_y$ represents the market-power distortion calculated as the inverse of the steady-state markup of the real wage over the marginal product of labor.[41] This result can be used in (51) to obtain (dropping the z_t^2 exogenous term)

$$S_t \simeq yU_c(1-\Phi_y)\left(\widehat{y}_t + \frac{1}{2}\left(\frac{\alpha+\theta_w^{-1}}{1-\alpha} + \frac{1}{\theta_p}\right) var_i\widehat{y}_t(i)\right)$$
$$+ \, yU_c(1-\Phi_y)\left(\frac{1}{2}(1+\gamma)(1-\alpha)^{-1}\widehat{y}_t^2 - \widehat{y}_t(1+\gamma)z_t\right). \quad (52)$$

Using (45a) for V_t and (52) for S_t, the social utility function (43) can be approximated (after some algebra) by

$$U_t \simeq yU_c\left(\Phi_y\widehat{y}_t - \frac{1}{2}\varpi^{-1}\widehat{y}_t^2 + \widehat{y}_t((1+\gamma)z_t + \chi_t)\right.$$
$$\left. -\frac{1}{2}\left(\frac{\alpha+\theta_w^{-1}}{1-\alpha} + \frac{1}{\theta_p}\right) var_i\widehat{y}_t(i)\right), \quad (53)$$

where $\varpi = \frac{1-\alpha}{\sigma(1-\alpha)+\alpha+\gamma}$ and the terms $\Phi_y\widehat{y}_t^2$, $\Phi_y\widehat{y}_t((1+\gamma)z_t+\chi_t)$, and $\Phi_y var_i\widehat{y}_t(i)$ were neglected as being of order higher than 2 as in Woodford (2003, 393–94). Let us define the relative-to-efficiency output gap as

$$\widetilde{y}_t - \widetilde{y}^* = (\widehat{y}_t - \widehat{\overline{y}}_t) - \widetilde{y}^*, \quad (54)$$

where, again following Woodford (2003, 395), $\widetilde{y}^*$ is the efficient level of the output gap obtained as the fractional difference in steady state between the level of output produced in a perfectly competitive economy and the level of potential output obtained in a monopolistically competitive economy, i.e., $\widetilde{y}^* = \log(\frac{y^*}{\overline{y}})$. For the FSW model described in the text, we have $\widetilde{y}^* = -\varpi\log(1-\Phi_y) \simeq \varpi\Phi_y$, where ϖ completely depends on the values of parameters regarding preferences and technology as defined above and $\Phi_y = \theta_p^{-1}$. From $(\widetilde{y}_t - \widetilde{y}^*)^2$ implied by (54) and $\widetilde{y}^* = \varpi\Phi_y$, the square output fluctuations, $\widehat{y}_t^2$, can be written as follows:

$$\widehat{y}_t^2 = (\widetilde{y}_t - \widetilde{y}^*)^2 + 2\widehat{y}_t\widehat{\overline{y}}_t - \widehat{\overline{y}}_t^2 - (\varpi\Phi_y)^2 + 2\varpi\Phi_y\widehat{y}_t - 2\varpi\Phi_y\widehat{\overline{y}}_t, \quad (55)$$

[41] Note that in the FSW model, $\Phi_y = \theta_p^{-1}$ and $1-\Phi_y = \frac{\theta_p - 1}{\theta_p}$.

where potential output fluctuations are $\widehat{\overline{y}}_t = \varpi((1+\gamma)z_t + \chi_t)$ as in equation (33) of the text. Inserting (55) into (53), we obtain the following (after dropping constant and exogenous terms):

$$U_t \simeq -\frac{yU_c}{2}\left(\varpi^{-1}(\widetilde{y}_t - \widetilde{y}^*)^2 + \left(\frac{\alpha+\theta_w^{-1}}{1-\alpha} + \frac{1}{\theta_p}\right) var_i\widehat{y}_t(i)\right). \quad (56)$$

By log-linearizing (48), the variance on differentiated output can be expressed in terms of price dispersion:

$$var_i\widehat{y}_t(i) = \theta_p^2 var_i\widehat{P}_t(i),$$

which leaves (56) as follows:

$$U_t \simeq -\frac{yU_c}{2}\left(\varpi^{-1}(\widetilde{y}_t - \widetilde{y}^*)^2 + \theta_p\left(1 + \frac{\theta_p\left(\alpha+\theta_w^{-1}\right)}{1-\alpha}\right) var_i\widehat{P}_t(i)\right).$$

Accordingly, the central bank intertemporal (social) utility function $\sum_{j=0}^{\infty}\beta^j E_t U_{t+j}$ becomes

$$\sum_{j=0}^{\infty}\beta^j E_t U_{t+j} \simeq -\frac{yU_c}{2}\sum_{j=0}^{\infty}\beta^j E_t\left(\varpi^{-1}(\widetilde{y}_{t+j} - \widetilde{y}^*)^2 \right.$$
$$\left. + \theta_p\left(1 + \frac{\theta_p\left(\alpha+\theta_w^{-1}\right)}{1-\alpha}\right) var_i\widehat{P}_{t+j}(i)\right). \quad (57)$$

Applying the results of Woodford (2003, 694–96), due to our Calvo-style sticky-price structure, we obtain

$$var_i\widehat{P}_t(i) \simeq \eta var_i\widehat{P}_{t-1}(i) + \frac{\eta}{1-\eta}\left(\pi_t^p\right)^2,$$

which implies

$$\sum_{j=0}^{\infty}\beta^j E_t var_i\widehat{P}_{t+j}(i) \simeq \frac{\eta}{(1-\eta)(1-\beta\eta)}\sum_{j=0}^{\infty}\beta^j E_t\left(\pi_{t+j}^p\right)^2. \quad (58)$$

Combining (57) and (58), we obtain

$$\sum_{j=0}^{\infty} \beta^j E_t U_{t+j} \simeq -\frac{yU_c}{2} \sum_{j=0}^{\infty} \beta^j E_t \left(\varpi^{-1} (\widetilde{y}_{t+j} - \widetilde{y}^*)^2 + \frac{\theta_p \left(1 + \frac{\theta_p \left(\alpha + \theta_w^{-1} \right)}{1-\alpha} \right) \eta}{(1-\eta)(1-\beta\eta)} \left(\pi_{t+j}^p \right)^2 \right),$$

which can be reorganized as follows:

$$\sum_{j=0}^{\infty} \beta^j E_t U_{t+j} \simeq -\Xi \sum_{j=0}^{\infty} \beta^j E_t \left(\left(\pi_{t+j}^p \right)^2 + \varpi^{-1} \frac{(1-\eta)(1-\beta\eta)}{\theta_p \left(1 + \frac{\theta_p \left(\alpha + \theta_w^{-1} \right)}{1-\alpha} \right) \eta} (\widetilde{y}_{t+j} - \widetilde{y}^*)^2 \right),$$

with $\Xi = \frac{yU_c}{2} \frac{\theta_p \left(1 + \frac{\theta_p (\alpha + \theta_w^{-1})}{1-\alpha} \right) \eta}{(1-\eta)(1-\beta\eta)}$. Our last result implies that the central bank loss function for period t is

$$L_t = \left(\pi_t^p \right)^2 + \lambda (\widetilde{y}_t - \widetilde{y}^*)^2,$$

with $\lambda = \frac{(1-\eta)(1-\beta\eta)(\frac{\alpha+\gamma}{1-\alpha} + \sigma)}{\eta \left(1 + \frac{\theta_p (\alpha + \theta_w^{-1})}{1-\alpha} \right) \theta_p}$ as defined in section 6 for the output-gap weight in the welfare-theoretic loss function of the FSW model.

Appendix 4. Welfare-Theoretic Targeting Rule in the FSW Model

Using the "timeless perspective" optimality criterion (Woodford 1999, 18), optimal monetary policy can be reached by computing the targeting rule that minimizes the intertemporal welfare-theoretic loss function subject to a set of equations describing the FSW model. Formally, it can be written as follows:

$$\underset{\pi_t^p, \widetilde{y}_t, u_t, \widehat{w}_t, \pi_t^w}{Min} \quad E_t \sum_{j=0}^{\infty} \beta^j L_{t+j}$$

subject to these all-time constraints:

$$\pi^p_{t+j} = \beta E_{t+j}\pi^p_{t+1+j} + \phi_1\left(\widehat{w}_{t+j} + \frac{\alpha}{1-\alpha}\widetilde{y}_{t+j}\right) - \frac{\phi_1}{\theta_w}u_{t+j} + \varkappa^{\pi^p}_{t+j}, \tag{59}$$

$$u_{t+j} = \frac{1}{\gamma}\widehat{w}_{t+j} - \phi_2\widetilde{y}_{t+j} + \varkappa^u_{t+j}, \tag{60}$$

$$\pi^w_{t+j} = \beta E_{t+j}\pi^w_{t+1+j} - \frac{\theta_p}{\theta_w(1-\alpha)}\left(\pi^p_{t+j} - \beta E_{t+j}\pi^p_{t+1+j}\right) - \phi_3 u_{t+j}, \tag{61}$$

$$\widehat{w}_{t+j} = \widehat{w}_{t-1+j} + \pi^w_{t+j} - \pi^p_{t+j}, \tag{62}$$

for $j = \ldots, -2, -1, 0, 1, 2, \ldots$. Equation (59) is the inflation equation (30) from the text written for period $t+j$ because it should be noted that $\phi_1 = \frac{(1-\eta)(1-\beta\eta)}{\eta\left(1+\frac{\theta_p(\alpha+\theta_w^{-1})}{1-\alpha}\right)}$, $\widehat{\psi}_{t+j} = \widehat{w}_{t+j} - \widehat{mpl}_{t+j} = \widehat{w}_{t+j} + \frac{\alpha}{1-\alpha}\widehat{y}_{t+j} = \widehat{w}_{t+j} + \frac{\alpha}{1-\alpha}(\widetilde{y}_{t+j} + \widehat{\overline{y}}_{t+j})$, and $\varkappa^{\pi^p}_{t+j} = \phi_1\frac{\alpha}{1-\alpha}\widehat{\overline{y}}_{t+j}$. Next, (60) relates the unemployment rate to real-wage fluctuations, the output gap, and the exogenous term $\varkappa^u_{t+j} = -\phi_2\widehat{\overline{y}}_{t+j} + z_{t+j} + \frac{1}{\gamma}\chi_{t+j}$ with $\phi_2 = \frac{\sigma(1-\alpha)+\gamma}{\gamma(1-\alpha)}$. Some algebra is involved in the determination of (60). In short, the unemployment rate (15) can be decomposed in $u_{t+j} = (\widehat{n}^s_{t+j} - \widehat{\overline{n}}_{t+j}) + (\widehat{\overline{n}}_{t+j} - \widehat{n}_{t+j})$, where the second term is $(\widehat{\overline{n}}_{t+j} - \widehat{n}_{t+j}) = -\frac{1}{1-\alpha}\widetilde{y}_{t+j}$. Using the log-linear version of the labor-supply curve (5) both in current and flexible-price observations, we find $\widehat{n}^s_{t+j} - \widehat{\overline{n}}_{t+j} = \frac{1}{\gamma}\widehat{w}_{t+j} - \frac{\sigma}{\gamma}\widetilde{y}_{t+j} + \varkappa^u_{t+j}$, which, substituted for the first term of the unemployment decomposition, yields (60). Next, the wage-inflation equation of the FSW model, equation (19), has been adapted to period $t+j$ to be written in the form of (61) with $\phi_3 = \frac{(1-\eta)(1-\beta\eta)}{\eta\theta_w}$. Finally, the auxiliary equation (62) was already introduced in the text as equation (40). For the optimal monetary policy conducted in period t, the central bank first-order conditions are

$$2\pi^p_t + \xi^{\pi^p}_t - \xi^{\pi^p}_{t-1} + \frac{\theta_p}{\theta_w(1-\alpha)}\xi^{\pi^w}_t - \frac{\theta_p}{\theta_w(1-\alpha)}\xi^{\pi^w}_{t-1} + \xi^w_t = 0, \tag{63}$$

$$2\lambda(\widetilde{y}_t - \widetilde{y}^*) - \phi_1\frac{\alpha}{1-\alpha}\xi^{\pi^p}_t + \phi_2\xi^u_t = 0, \tag{64}$$

$$\frac{\phi_1}{\theta_w}\xi_t^{\pi^p} + \xi_t^u + \phi_3\xi_t^{\pi^w} = 0, \quad (65)$$

$$-\phi_1\xi_t^{\pi^p} - \frac{1}{\gamma}\xi_t^u + \xi_t^w - \beta E_t\xi_{t+1}^w = 0, \quad (66)$$

$$\xi_t^{\pi^w} - \xi_{t-1}^{\pi^w} - \xi_t^w = 0, \quad (67)$$

where $\xi_t^{\pi^p}$, ξ_t^u, $\xi_t^{\pi^w}$, and ξ_t^w are the Lagrange multipliers associated with constraints (59)–(62) in period t. The targeting rule that defines the optimal monetary policy consists of equations (59)–(67), which provide solution paths for the nine endogenous variables π_t^p, π_t^w, $\widehat{w}_t$, u_t, $\widetilde{y}_t$, $\xi_t^{\pi^p}$, ξ_t^u, $\xi_t^{\pi^w}$, and ξ_t^w. The IS curve (31) together with the output-gap definition, $\widetilde{y}_t = \widehat{y}_t - \widehat{\overline{y}}_t$, can be added to the system to determine current output and the required move on the nominal interest rate, $\widehat{y}_t$ and R_t.

Appendix 5. Welfare-Theoretic Targeting Rule in the EHL Model

Using the "timeless perspective" optimality criterion (Woodford 1999, 18), optimal monetary policy can be reached by computing the targeting rule that minimizes the intertemporal welfare-theoretic loss function subject to a set of equations describing the EHL model. Formally, it can be written as follows:

$$\underset{\pi_t^p, \widetilde{y}_t, \widehat{w}_t, \pi_t^w}{Min} \quad E_t \sum_{j=0}^{\infty} \beta^j \mathcal{L}_{t+j}$$

subject to these all-time constraints:

$$\pi_{t+j}^p = \beta E_{t+j}\pi_{t+1+j}^p + \varphi_1\left(\widehat{w}_{t+j} + \frac{\alpha}{1-\alpha}\widetilde{y}_{t+j}\right) + \tau_{t+j}^{\pi^p}, \quad (68)$$

$$\pi_{t+j}^w = \beta E_{t+j}\pi_{t+1+j}^w + \varphi_2\left(\left(\frac{\gamma}{1-\alpha} + \sigma\right)\widetilde{y}_{t+j} - \widehat{w}_{t+j}\right) + \tau_{t+j}^{\pi^w}, \quad (69)$$

$$\widehat{w}_{t+j} = \widehat{w}_{t-1+j} + \pi_{t+j}^w - \pi_{t+j}^p, \quad (70)$$

for $j = \ldots, -2, -1, 0, 1, 2, \ldots$. Equation (68) is the New Keynesian Phillips curve of the EHL model, equation (36b), written in terms of π^p_{t+j} because $\varphi_1 = \frac{(1-\eta_p)(1-\beta\eta_p)}{\eta_p(1+\frac{\theta_p}{1-\alpha})}$ and $\tau^{\pi^p}_{t+j} = \varphi_1 \frac{\alpha}{1-\alpha}\widehat{\overline{y}}_{t+j}$. The wage-inflation equation of the EHL model, equation (37b), has been adapted to period $t + j$ to be written in the form of (69), with $\varphi_2 = \frac{(1-\eta_w)(1-\beta\eta_w)}{\eta_w(1+\gamma\widetilde{\theta}_w)}$ and $\varkappa^{\pi^w}_{t+j} = \varphi_2((\frac{\gamma}{1-\alpha}+\sigma)\widehat{\overline{y}}_{t+j} - \frac{\gamma}{1-\alpha}z_{t+j} - \chi_{t+j})$. Finally, the auxiliary equation (70) was already introduced above in (30) for period t. For the optimal monetary policy in period t, the central bank first-order conditions are

$$2\lambda_{\pi^p}\pi^p_t + \xi^{\pi^p}_t - \xi^{\pi^p}_{t-1} + \xi^w_t = 0, \tag{71}$$

$$2\lambda_{\widetilde{y}}(\widetilde{y}_t - \widetilde{y}^*) - \varphi_1 \frac{\alpha}{1-\alpha}\xi^{\pi^p}_t - \varphi_2\left(\frac{\gamma}{1-\alpha} + \sigma\right)\xi^{\pi^w}_t = 0, \tag{72}$$

$$-\varphi_1\xi^{\pi^p}_t + \varphi_2\xi^{\pi^w}_t + \xi^w_t - \beta E_t\xi^w_{t+1} = 0, \tag{73}$$

$$\xi^{\pi^w}_t - \xi^{\pi^w}_{t-1} - \xi^w_t = 0, \tag{74}$$

where $\xi^{\pi^p}_t$, $\xi^{\pi^w}_t$, and ξ^w_t are the Lagrange multipliers associated with constraints (68)–(70) in period t. The targeting rule that defines the optimal monetary policy in the EHL model consists of equations (68)–(74), which provide solution paths for the seven endogenous variables π^p_t, π^w_t, $\widehat{w}_t$, $\widetilde{y}_t$, $\xi^{\pi^p}_t$, $\xi^{\pi^w}_t$, and ξ^w_t. The IS curve (31) together with the output-gap definition, $\widetilde{y}_t = \widehat{y}_t - \widehat{\overline{y}}_t$, can be added to the system to determine current output and the required move on the nominal interest rate, $\widehat{y}_t$ and R_t.

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