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Expectations as a Source of Macroeconomic Persistence: Evidence from Survey Expectations in Dynamic Macro Models

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Abstract:

This paper finds that persistence intrinsic to expectations may explain a sizable fraction of the persistence in aggregate macroeconomic time series. The paper endogenizes survey expectations in an array of standard macroeconomic relationships and in a DSGE model. The use of survey measures of expectations—for near-term inflation, long-term inflation, near-term and long-term unemployment, and short-term interest rates—improves performance along a variety of dimensions. Survey expectations exhibit strong correlations to key macroeconomic variables. Using a minimal set of assumptions, those correlations may be given a structural interpretation in a DSGE context. Including survey expectations helps to identify key slope parameters in standard relationships, and nearly eliminates the need for lagged dependent variables in structural models that is often motivated by indexation for prices and habit formation for consumption. Including survey expectations also obviates the need for autocorrelated structural shocks in the key equations. The paper discusses the modeling complications that arise once the rational expectations assumption is abandoned, and proposes methods for endogenizing survey expectations in a general equilibrium macro model. Overall, the results suggest that much of the persistence in aggregate data is better accounted for by slow-moving expectations, rather than by habits, indexation and autocorrelated structural shocks.

JEL Classifications: E32, E52

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1. Introduction

Over the past decade, macroeconomists have converged on dynamic stochastic general equilibrium (DSGE) models with rational expectations as the standard for macroeconomic modeling. All these models feature a prominent role for expectations—in wage- and price-setting, consumption and investment spending, and asset prices—that most economists agree is essential to constructing a realistic depiction of economic behavior. A number of authors have documented the empirical success of these DSGE models and have suggested that they should serve as useful inputs to formulating monetary and fiscal policy (see, for example, Smets and Wouters 2003 and Christiano, Eichenbaum, and Evans 2005). In the aftermath of the Great Recession, many economists have proposed augmentations to these DSGE models that could guide policies designed to mitigate bouts of financial instability.

In most cases the expectations in these models are assumed to be rational, in the sense that all agents' expectations are assumed to equal the mathematical expectations implied by the DSGE model. Yet a growing body of work suggests that simple DSGE models with rational expectations demonstrate significant counterfactual implications (see for example Estrella and Fuhrer 2002, Rudd and Whelan 2005).[1] Partly in response to such criticisms, a number of authors have proposed augmentations to earlier vintages of DSGE models that better allow the models to match many of the key moments in the data. The additions of habit formation, price indexation, adjustment costs, and serially correlated shocks all fall into this category. It is important to note that the microeconomic evidence in favor of habit formation (see, for example, Dynan 2000) is mixed, and there is virtually no evidence in microeconomic price data of indexation. Direct evidence on the time series properties of shocks is necessarily limited. While one can easily imagine that some shocks exhibit persistence over time (shocks to energy prices are a leading example in recent years),

[1] Much earlier work emphasized the unrealistic information assumptions implied by the rational expectations hypothesis. For an early and important example, see Friedman (1979).

modelers may wish to strike a better balance between allowing for some persistence in shocks and attributing too much of the business cycle fluctuations in macroeconomic data to the time series properties of unobservable shocks.

All of these augmentations, however, are conducted using the rational expectations paradigm. This paper investigates the extent to which a change in the expectations assumption can substitute for these augmentations, thus resulting in a model that retains many of the underlying structures that have been developed in recent years, but without some of the "bells and whistles" (or, employing an astronomical analogy, "epicycles") that have allowed the models to meet formidable empirical challenges.

To be sure, a large literature explores alternative expectations schema, in some cases employing survey data to help identify expectations mechanisms. An early example of such a strategy is Roberts (1997), which uses survey expectations in an estimated New Keynesian Phillips curve.[2] This paper also complements recent research by Fuster, Hebert, and Laibson (2012), who examine the implications of more realistic (or "natural") expectation formation. Carroll (2003) explores the "epidemiological" transmission of expectations from professional (SPF) to household (Michigan) surveys, using the aggregate data from both of these sources. There is also a large theoretical and empirical literature on adaptive learning (see Adam (2005), the many papers of Evans and Honkapohja and their 2001 book, Milani (2007), Orphanides and Williams (2005), and Slobodyan and Wouters (2012)). Milani (2007) shows that the introduction of adaptive learning significantly reduces the dependence of a particular DSGE model on habit formation and price indexation in explaining the persistence of macroeconomic time series. Slobodyan and Wouters (2012) find a notable reduction in the persistence of the estimated shocks that drive wages and

[2] A very early paper by Lahiri and Lee (1979) explores these issues using a different set of methodologies and a (necessarily) earlier sample period. More recently, Piazzesi and Schneider (2009) have examined the usefulness of survey expectations in affine term structure models.

prices; they also note that the expectations based on the "small forecasting models" in their paper bear a close resemblance to survey expectations. Another line of research that lies close to the tack taken in this paper employs survey data as observations that help in estimating rational expectations and learning models. A recent and insightful example is Molnár and Ormeño (forthcoming), which uses the survey data to add moment restrictions to both rational expectations and adaptive learning models. They find that the addition of these restrictions improves the performance of the learning model relative to rational expectations.

This paper takes a different tack, instead using survey expectations directly as the expectations in the model. Making a minimal set of assumptions about the properties of survey expectations allows the survey data to assume the role of the model's expectations, rather than to augment estimation and identification of more complex expectations structures, such as learning and sticky information (see Mankiw and Reis 2002). The assumptions that are employed in the paper follow Adam and Padula (2011) and Branch and McGough (2009), allow one to aggregate across heterogeneous agents, to ignore higher-order expectations, to pass expectations through linear operators, and importantly to allow surveys to conform to the law of iterated expectations. With these assumptions, one can derive models that embed survey expectations, and that are reasonably well-approximated by the underlying log-linearized relationships embodied in standard DSGE models. Thus output depends in the conventional manner on expected output and real interest rates; inflation depends on expected inflation and real output or marginal cost; and short-term interest rates are set by a monetary authority according to a forward-looking policy rule that depends on expected inflation and output relative to their targets.

Using this approach, the paper develops evidence that the systematic use of survey expectations—one way of incorporating measured expectations, rather than assuming rational expectations—offers a number of advantages over the rational expectations models. The identification of key parameters is improved, and the need for macroeconomic "epicycles" such as correlated shocks and pseudo-structural features that add lagged endogenous variables to the model is obviated. The empirical success of the survey-based DSGE model is encouraging.

Surveys now provide rather extensive data on the forecasts and expectations formed by agents in the economy. While the incentives to devote resources to expectation-formation are questionable in some surveys, for the respondents to the SPF employed in this paper, forecasting is a primary business line for the survey participants, so presumably the incentives are strong for devoting significant resources to forecasting.[3] The paper will not test the extent to which survey expectations may be considered "rational" in the statistical senses of unbiased and efficient; many authors have done so in previous work (see Batchelor 1986, Bryan and Gavin 1986, Mehra 2002, Thomas 1999, and Adam and Padula 2011). Instead, this paper will take the survey expectations as given, despite the possibility that such expectations may be characterized by irrationality.

Of course, such a departure from rationality comes at a cost. The beauty of the rational expectations paradigm is that it instantly answers many questions about how expectations evolve. If one is willing to specify a model, one simultaneously has specified the expectations that are consistent with the model. That beauty is lost with the introduction of survey expectations, as one can no longer "solve out" expectations in the simple way that has become standard in the DSGE literature. Choosing to use survey expectations necessitates the use of theory-based approximations and empirically-motivated compromises, which will be described in more detail below.

Recognizing these tradeoffs, the paper concludes that the move to employing survey-based measures of expectations represents a viable and potentially useful direction for macroeconomic modeling. Section 2 provides a simple theory example that illustrates the challenges in departing from rational expectations, and then details the approach to expectations formation that is pursued

[3] Other incentives may also influence the behavior of economic forecasters, such as the desire to distinguish one's forecast from other forecasters in order to gain market share; see, for example Batchelor and Dua 1990.

in this paper, itemizing the assumptions necessary to embed survey expectations in a dynamic macro model. Section 3 develops a DSGE model that employs an array of survey expectations measures, is consistent with the core of extant theory, and employs the assumptions of section 2 to address the theoretical difficulties inherent in departing from rational expectations. Section 4 presents some single-equation evidence that suggests that a variety of survey expectations measures may be helpful in key elements of macroeconomic models. Section 5 presents an array of empirical results from system estimation of the model developed in section 3, along with a variety of tests to assess the relative contributions to explaining aggregate dynamics from persistence in the expectations process, versus persistence arising from habits, indexation and autocorrelated shocks. Section 6 concludes.

2. Expectations Formation

Examining a simple rational expectations model will help to illustrate the issues that arise in using survey expectations to model macroeconomic behavior. The example will also clarify some of the mechanics surrounding the use of rational expectations, which may be helpful for those who do not routinely solve rational expectations models.

Consider a two-equation model that describes the evolution of inflation (π) in a manner similar to that in Calvo (1983) or Rotemberg (1983). A simple process for output (y) closes the model:

$$\begin{aligned} \pi_t &= \beta E_t \pi_{t+1} + \gamma y_t \\ y_t &= \alpha y_{t-1} + \varepsilon_t \end{aligned} \quad . \tag{2.1}$$

One can iterate equation (2.1) forward successively, substituting in future expectations of y_{t+i} as follows:

$$\begin{aligned} \pi_t &= \gamma y_t + \gamma\beta E_t y_{t+1} + \gamma\beta^2 E_t y_{t+2} + \ldots + \varepsilon_t \\ &= \gamma y_t \sum_{i=0}^{\infty} (\alpha\beta)^i + \varepsilon_t \end{aligned} \quad , \tag{2.2}$$

using the definition of *y* given in the second equation in (2.1) to substitute for all occurrences of $E_t y_{t+i}$ as $\alpha^i y_t$, and assuming that $E_t(\varepsilon_{t+i}) = 0 \;\; \forall i > 0$. Alternatively, one can solve out for the unobserved quantity $E_t \pi_{t+1}$ as a function of observed inflation and output to obtain a constrained version of equation (2.2) that is based only on observables, but maintains this equations's underlying structural form:[4]

$$\pi_t = Bz_t + \gamma y_t ; z_t \equiv \begin{bmatrix} \pi_t \\ y_t \end{bmatrix}. \tag{2.3}$$

The symbol *B* represents a row from the coefficient matrix that defines the model's restricted reduced-form solution. In this case, $B = \begin{bmatrix} 0 & b_2 \end{bmatrix}$ and b_2 depends in a relatively straightforward way on the underlying structural parameters $[\beta, \gamma, \alpha]$.

Now consider the same model with survey, rather than rational, expectations. The survey expectation of inflation for period *t+1* made in period *t* is denoted as $\pi^S_{t+1,t}$

$$\begin{aligned} \pi_t &= \pi^S_{t+1,t} + \gamma y_t \\ y_t &= \alpha y_{t-1} + \varepsilon_t \end{aligned} . \tag{2.4}$$

In this case, it is less clear how to solve the model. On the one hand, because the survey expectations are observable, one need not "solve out" the expectations in equation (2.4) in order to pin down inflation and output. However, the model is not fully closed, as no process for $\pi^S_{t+1,t}$ is necessarily implied by equation (2.4). Should one wish to simulate the model forward in time from arbitrary initial conditions, or examine its behavior under alternative policy assumptions, one would need to specify how the survey expectations would evolve over time.

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[4] The methods of Blanchard and Kahn (1980) and Anderson and Moore (1983) implement such solutions.

This paper draws on the work of Adam and Padula (2011) and Branch and McGough (2009), using a minimal set of assumptions for expectation formation that allows one to incorporate (potentially) heterogeneous expectations in log-linearized Euler equations for inflation and output. One key assumption is that the survey expectations may be consistently iterated forward, as assumed both in Adam and Padula (2011) and in Branch and McGough (2009). The assumption summarized in equation (2.5) below ensures that individual agents expect no predictable revision in their own or in others' forecasts for any period in the forecast horizon. For the forecast operator *F*, the ith forecaster's expected revision of forecaster *h* from period *t* to *t+1* is 0:

$$\text{Assumption 1: } F_t^i[F_{t+1}^h[X_{t+s}]-F_t^h[X_{t+s}]]=0 \ \forall \ i, h, s>0 \ . \tag{2.5}$$

If this assumption holds, then the subjective forecast of the aggregate inflation rate is a sufficient statistic for all the agents' forecasts—"higher-order" expectations of other forecasters' forecasts are irrelevant. Although it is common in the literature, this is a fairly strong assumption. A somewhat weaker assumption, embedded in equation (2.5) when *i=h,* asserts that each forecaster expects no revision in her own forecast ($F_t^i[F_{t+1}^i[X_{t+s}]-F_t^i[X_{t+s}]]=0, s>0$), which in turn implies that expectations obey the law of iterated expectations. This assumption is employed in the derivations below, without implying that expectations are rational in the conventional sense. Allowing for the law of iterated expectations to hold gets us part way towards a solution: consider the first step of such an iteration displayed below:

$$\begin{aligned} \pi_t &= \pi_{t+1,t}^S + \gamma y_t + \varepsilon_t \\ \pi_{t+1,t}^S &= \pi_{t+2,t}^S + \gamma y_{t+1,t}^S \, . \end{aligned} \tag{2.6}$$

But now one needs to know the survey-based expectation for *output* in period *t+1*. If this data is available, that solves the problem for the first step of the iteration, but it should be obvious by now that the iteration continues, and one cannot use the output process in equation (2.4) to quickly solve for all future output expectations (doing so would return us to the rational expectations solution). At some point, one will run out of survey expectations for farther-forward observations, and the question of how to close the model thus remains.

One can rewrite equation (2.6) to partition the forward iteration of the equation into two components (ignoring for the moment the shock ε_t):

$$\pi_t = \gamma y_t + \Gamma[y_{t+1}, y_{t+2}, \ldots] \ , \tag{2.7}$$

where the function $\Gamma[.]$ is a (presumably discounted) sum of future output terms. The strategy employed in this paper is to use the long-run expectations of key variables to proxy for $\Gamma[.]$. For the Phillips curve, by assumption, the basic structure of the model implies that long-run inflation expectations should embody expectations for output well out into the future—however those expectations are formed. For the consumption Euler equation, longer-run output expectations should similarly embody expectations for real interest rates (denoted ρ_{t+i} below) out into the future. Thus our second key assumption for expectations formation is:

$$\text{Assumption 2: } \Gamma[y_{t+1}, y_{t+2}, \ldots] \propto \Pi^S_{LR,t}; \Gamma[\rho_{t+1}, \rho_{t+2}, \ldots] \propto y^S_{LR,t} \ , \tag{2.8}$$

where $\Pi^S_{LR,t}$ and $y^S_{LR,t}$ denote the long-run survey expectations for inflation and output. In the case of inflation expectations, this would imply that (2.7) becomes

$$\pi^S_{t+1,t} \approx \gamma y_t + b\Pi^S_{LR,t} \tag{2.9}$$

The modeling exercise presented in section 4 below thus employs an expectations process that implies that (a) in the short-run, survey expectations may be consistently iterated forward, obeying the law of iterated expectations; (b) higher-order expectations do not enter into the expectations process; (c) long-run expectations of inflation (output) proxy for the longer-horizon expectations of output (real interest rates) in the Phillips and IS curves; (d) as in Branch and McGough (2009), at long horizons, survey expectations converge towards the long-run rational expectations equilibrium for the model[5]; consequently, the long-run expectation implied by the surveys will equal the model's steady state value for that variable.

$$\text{Assumption 3: } x_{\infty}^{S} \equiv x_{Model}^{SS} \tag{2.10}$$

Implicitly, in adopting the standard New Keynesian Euler equations for inflation and output, we also adopt Branch and McGough's (2009) assumptions A.1, A.3 and A.4, which imply that the expectation of observable realizations equal the realizations (A.1), and that the expectations operator may be passed through simple linear operators (A.3 and A.4).[6] The details of this strategy will be more fully discussed in section 3, and will ultimately be guided by empirical considerations as well as theoretical purity.

Finally, the expectations process allows for the possibility of "intrinsic inertia" in expectations, although the degree of such inertia will be estimated, and estimated jointly with other parameters that index model features that could also impart inertia to inflation and output. In particular, short-run expectations may adjust gradually, rather than immediately, to the expectations implied by the forward-iterated Euler equations. Using the one period-ahead inflation expectation equation (2.9) as an example, the partial-adjustment equation is:

$$\text{Assumption 4: } \pi_{t+1,t}^{S} = \mu^{\pi}[b\Pi_{LR,t}^{S} + \gamma y_t] + (1-\mu^{\pi})(\pi_{t,t-1}^{S}) \ , \tag{2.11}$$

where $(1-\mu^{\pi})$ indexes the speed of adjustment of expectations toward the expectations proxy in equation (2.9). Thus short-run expectations will be anchored by the expectations implied by equation(2.9), but may move toward the anchor gradually, with a speed determined by partial adjustment coefficient μ . Such inertial expectations behavior is strongly validated by the micro-data for both the SPF and for the University of Michigan survey of consumers, which exhibit a strong propensity for individual forecasters to anchor their forecasts to lagged central tendencies of forecasts, see Fuhrer (2015).

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[5] See assumption A.2 in Branch and McGough, p. 1038.

[6] That is, operations such as E(ax)= aE(x), and E(x+y+z) = E(x)+E(y)+E(z), will hold in the survey expectations.

To summarize, we construct an expectations mechanism that comprises Assumptions 1-4, which in turn build on the assumptions of Adam and Padula (2011) and Branch and McGough (2009). These assumptions jointly allow the law of iterated expectations to hold, allow us to employ aggregate Euler equations for inflation and output, and flexibly allow for some degree of sluggish adjustment of expectations to fundamentals. As we will see below, the sluggish expectation adjustment mechanism specified in (2.11) is essential in capturing the dynamics of inflation and output expectations and realizations.

3. A Structural DSGE Model with Ubiquitous Survey Expectations

Price-Setting

The survey-based model for price-setting follows closely the expectations strategy described in section 2. In motivating the model, we take a couple of steps back relative to the most recent DSGE models that include capital and wages, and begin with simpler formulations. This approach is adapted partly for simplicity and partly because many additions to the earlier models were made in response to the deficiencies observed in those models. Part of the goal of this paper is to determine to what extent those deficiencies arose from the assumption of rational expectations.

Under the standard assumptions underlying the Calvo formulation of sticky prices, augmented by the expectations assumptions of Adam and Padula (2011) and Branch and McGough (2009) as described in section 2 above, the behavior of inflation with survey expectations and without indexation will be defined by the difference equation,

$$\pi_t = \beta\pi^S_{t+1,t} + \lambda mc_t \,. \tag{3.1}$$

For a reasonable set of assumptions, one can show that marginal cost will be proportional to either the output gap or the unemployment gap, so that we can equivalently write equation (3.1) as[7]

$$\pi_t = \beta\pi^S_{t+1,t} - \pi^u(U_t - U^*_t)\,, \tag{3.2}$$

[7] See Blanchard and Galí (2010) for a derivation of a New Keynesian Philips curve with unemployment.

where π^u is a function of the standard Calvo parameters underlying λ, as well as the parameter on hours in the utility function. Our measure of inflation is the overall or "headline" CPI, which we choose because the longest-available long-dated (10-year) survey expectations measure from the SPF reports forecasts for the CPI. In order to aid in identifying key macro parameters, we allow for the independent effects of food and energy prices on the CPI as "supply shifters," thus augmenting the simple Phillips curve as follows,[8]

$$\pi_t = \beta\pi^S_{t+1,t} - \pi^u(U_t - U^*_t) + w^e dp^e_t + w^f dp^f_t \ . \qquad (3.3)$$

Apart from the inclusion of food and energy price shocks, the only substantive difference between this equation and the standard simple DSGE models is the use of survey, rather than rational, expectations.

To endogenize inflation expectations, we proceed as in section 2, approximating the long sequence of expected unemployment gaps with the SPF measure of the 10-year average expected inflation rate, as this measure should embody—according to the model's underlying logic—the appropriate sequence of short-term expectations, in a sense performing the forward iteration for us. Thus we posit the inflation expectations equation,

$$\pi^S_{t+1,t} = A^{\pi e}\Pi^S_{LR,t} - \pi^u(U^S_{t+1,t} - U^*_{t+1}) \ . \qquad (3.4)$$

Of course, to completely close the model, we will need to solve for the survey expectations for unemployment in subsequent periods. We will tackle this issue when we discuss the IS curve below.

As suggested in section 2, we allow for the possibility that short-run expectations will adjust gradually towards the long-run sequence of expectations implied by equation (3.4), and thus allow for an error-correction equation much like (2.11)

[8] A measure of marginal cost that captured all marginal input costs would presumably include such cost shocks, but in the simple model here, in which marginal cost is proportional to employment, one cannot assume that this will be the case.

$$\pi_{t+1,t}^{S} = \mu^{\pi}[A^{\pi e}\Pi_{LR,t}^{S} - \pi^{u}(U_{t+1,t}^{S} - U_{t+1}^{*})] + (1-\mu^{\pi})(\pi_{t,t-1}^{S}) \ . \tag{3.5}$$

where the parameter $(1-\mu^{\pi})$ indexes the speed of adjustment. Note that this partial adjustment formulation implicitly introduces lagged inflation expectations into the determination of inflation and inflation expectations; equivalently, it builds some "intrinsic persistence" into the inflation *expectations* process without introducing lags of realized inflation.[9] We will examine the importance of this partial adjustment mechanism in section 5, specifically by testing the data's ability to distinguish between the influence of lagged expected versus lagged realized inflation.

Energy and food prices, which enter equation (3.3), are assumed to follow simple AR processes in log changes:

$$dp_t^e = a^e dp_{t-1}^e; dp_t^f = a^f dp_{t-1}^f \tag{3.6}$$

Long-run inflation expectations are taken as a proxy for the central bank's current inflation goal, which varies over time. In this model, the current inflation goal (and long-run inflation expectations) are assumed ultimately to converge to the fixed long-run central bank inflation target π^*. The inflation goal can deviate from its long-run target with some persistence, which we model via a partial adjustment equation with parameter γ

$$\Pi_{LR,t}^{S} = \gamma\Pi_{LR,t-1}^{S} + (1-\gamma)\pi^{*} . \tag{3.7}$$

OLS estimates of equation (3.7) imply that γ has a value of 0.95, and this value is used throughout the remainder of the paper.[10] Note that we choose this slow-moving autoregressive process for the long-run inflation rate because it is a simple way of endogenizing a time-varying inflation goal, in a manner consistent with Assumption 3 of section 2, i.e. that in the long run, expectations converge to the model's steady state. The AR(1) captures the timeseries pattern of long-run inflation

[9] See Fuhrer (2006) for a definition of the concepts of "intrinsic" and "inherited" persistence in the context of New Keynesian Phillips curves.

[10] For a more complete model of the time-varying inflation goal, or "shiftint endpoint," see Kozicki and Tinsley (2012).

expectations, but does not impute deeper behavioral reasons for movements in the central bank's target; neither does this specification impose any significant restrictions on the rest of the model.

The incorporation of long-run expectations is in part consistent with the "trend inflation model" of Cogley and Sbordone (2008).[11] In those models, accounting for trend inflation often obviates the need for indexation. However, Fuhrer (2011) shows that this conclusion is very sensitive to estimation method. One can see this from an examination of the differences in the sample autocorrelations for raw and detrended inflation (using Cogley and Sbordone's (2008) trend), which are quite small—the first autocorrelation for inflation is 0.89 versus 0.81 in their longest sample.[12] We will show below that anchoring short-run expectations on long-run expectations does not come close to accounting for all of the dynamics in inflation.

IS Curve

Underlying the IS curves in most DSGE models is the simple life-cycle model of consumption, which under rational expectations and reasonable assumptions about preferences implies a linear approximation to the first-order conditions of the form,

$$c_t = \beta E_t c_{t+1} - \sigma(\rho_t - \overline{\rho}), \tag{3.8}$$

where ρ_t and $\overline{\rho}$ are a real rate of interest and the long-run equilibrium value of that rate, respectively. The approach used in many simple models in the literature is adopted here: it is assumed that capital investment is either absent or proportional to consumption, and government spending is fixed, so that equation (3.8) may equivalently be written as an output equation by simply substituting y_t for c_t, where y_t is understood to be the deviation of output from its equilibrium (perhaps flex-price) level. Finally, as discussed in the preceding subsection on price-setting, for simple production functions in which there is no difference between the intensive and the extensive

[11] Cogley and Sbordone (2008) are the first to consider so-called models of "trend inflation," in which inflation in the Calvo model is expressed as a deviation from trend inflation.
[12] See Fuhrer (2011) Table 11, p. 463.

margin of labor (i.e. we abstract from the difference between hours and employment), and in which capital input is fixed or absent, output and employment are proportional. One can thus substitute the unemployment gap for the output gap to arrive at an unemployment-based IS curve[13]:

$$U_t - U_t^* = u^{ue}(E_t U_{t+1} - U_{t+1}^*) + u^{\rho}(\rho_t - \overline{\rho}) \ . \tag{3.9}$$

The real interest rate here is defined as the difference between the nominal interest rate and the inflation expectation from the SPF.[14] Finally, dropping the rational expectations assumption and substituting survey expectations, and following the results in Branch and McGough (2009), we obtain a first-order difference equation in the unemployment gap and the real interest rate,

$$U_t - U_t^* = u^{ue}(U_{t+1,t}^S - U_{t+1}^*) + u^{\rho}(\rho_t - \overline{\rho}) . \tag{3.10}$$

As with the price equation, this equation is complete given observations on the one-period ahead survey expectations for unemployment, which are collected in the SPF. However, in order to close the model, we need to posit a process for the unemployment expectation, following the assumptions detailed in section 2. Thus we link the one-period-ahead inflation expectation to a long-term (e.g. ten-year average) unemployment expectation, which implicitly captures a sequence of short-run expectations of future real interest rates:

$$U_{t+1,t}^S - U_{t+1}^* = u^{ue}(U_{LR,t}^S - U_t^*) + u^{\rho}(\rho_{t+1} - \overline{\rho}) . \tag{3.11}$$

While the SPF does not collect such a variable on a consistent basis over a long sample, the Blue Chip forecast survey has done so since 1984.[15]

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[13] We could use the more standard IS curve, which is expressed in terms of an output gap. But because it is more difficult to construct a measure of equilibrium output that is consistent with the SPF forecasts of GDP through the years, given the changes in methodology and base years, we choose the form that is expressed in terms of the unemployment gap. In contrast to GDP, the civilian unemployment rate concept has remained relatively stable over the years. By using the CBO estimate of the equilibrium unemployment rate, we have left ambiguous how close our equilibrium unemployment measure is to the flex-price equilibrium implied by the model.

[14] Standard theory would suggest the one-period real interest rate should appear in the IS curve. A discussion of the choice of the interest rate in the IS curve that balances theoretical rigor and empirical identification appears in section 4.

[15] The 10-year average forecast is available beginning in March 1984, and is collected twice yearly, in March and October, thereafter. The use of this variable restricts the sample relative to the SPF, for which all observations are available by late

Once again we allow for data-determined partial adjustment towards the approximation to the sequence of expectations implied by equation (3.11). Consistent with the approach outlined in equation (2.11) of section 2, equation (3.12) specifies the partial adjustment of the short-run unemployment expectation to the longer-run expectations,[16]

$$U^S_{t+1,t} - U^*_{t+1} = \mu^U [u^{ue}(U^S_{LR,t} - U^*_t) + u^\rho(\rho_t - \overline{\rho})] + (1-\mu^U)(U^S_{t,t-1} - U^*_{t-1}), \tag{3.12}$$

where as in the specification for inflation expectations, the parameter $(1-\mu^U)$ indexes the speed of adjustment.[17]

Because the steady state for the unemployment gap should be zero in this simple model, we close the model by assuming that long-run unemployment expectations deviate temporarily from zero, in a manner parallel to long-run inflation expectations as defined in equation (3.7). That is,

$$U^S_{LR,t} - U^*_t = \gamma^U (U^S_{LR,t-1} - U^*_{t-1}) \ . \tag{3.13}$$

This approach guarantees that the long-run survey expectations will converge to the appropriate steady state for the unemployment gap (zero) in the long run. OLS estimates of equation (3.13) imply that γ^U has a value of 0.94, and this value is imposed throughout the remainder of the paper.

Interest Rates

In most DSGE models, the appropriate real rate of interest for the IS curve is the short-term risk-free real rate of interest. This formulation implies that, with rational expectations, real activity will implicitly depend on the long-term real interest rate (achieved by iterating forward the Euler equation into the infinite future). The standard definition of the short-term real interest rate ρ_t is

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1982. See the data appendix for details on the construction of the quarterly series for the 10-year average unemployment expectation.

[16] Initial estimation attempts that do not allow for partial adjustment experience more difficulty in identifying the key slope parameters in the IS relation, and also leave considerable unexplained variation in the unemployment expectations series.

[17] As with inflation, the difference between the autocorrelations of the raw unemployment data and the unemployment gap is modest: the first autocorrelation coefficient for the raw data is 0.94; the corresponding AR coefficient for the gap is 0.93.

just the difference between the current short-term policy rate and the one-period-ahead expected inflation rate,

$$\rho_t \equiv i_t - \pi^S_{t+1,t} \ . \tag{3.14}$$

We specify the policy rule that defines the short-term interest rate i_t. We take the short-term interest rate to be the central bank's policy rate, and develop a forward-looking policy rule that employs survey expectations of inflation, long-term inflation, and unemployment. First, define the deviation of the federal funds rate, $\tilde{i}_t$ from its long-run equilibrium as

$$\tilde{i}_t \equiv i_t - (\Pi^S_{L,t} + \overline{\rho}) \ . \tag{3.15}$$

We can then write a forward-looking policy rule in the policy rate deviation, allowing for interest-rate smoothing

$$\tilde{i}_t = a\tilde{i}_{t-1} + (1-a)[i^{\pi}(\pi^S_{t+1,t} - \Pi^S_{L,t}) - i^u(U^S_{t+1,t} - U^*_{t+1})] \ . \tag{3.16}$$

Note that in the model whenever we allow rational expectations to play a role, these expectations will of course take into account the role that survey expectations play in the short-run evolution of the key variables. This assumption will necessarily change the way in which rational expectations act in the model, as compared to standard DSGE models in which all expectations are rational.

For reasons discussed below, we also consider a slightly longer-term real interest rate in the IS block, defined as the difference between the SPF expectations for the one-year-ahead Treasury bill rate and the one-year-ahead inflation rate, or

$$\rho_t \equiv i^S_{1Y,t} - \pi^S_{1Y,t} \tag{3.17}$$

This compromise retains the spirit of the one-period trade-off implied by the consumption Euler equation and yields considerably better empirical performance, as will be shown below.[18] Of course,

[18] In a much earlier paper, Fuhrer and Moore (1995) showed that a reduced-form IS curve that depends explicitly on a longer-term real interest rate achieves some empirical success. See also Fuhrer and Rudebusch (2004) for a discussion of identification of the IS curve.

the one-year-ahead SPF forecast for the three-month Treasury bill, $i^S_{1Y,t}$ requires an internally-consistent definition to close the model. We assume that the forecasts implied by the policy rule for the federal funds rate will provide a reasonable approximation to the SPF's forecasts of the three-month Treasury bill rate over the next year. Thus the model's forecast of the average short rate over the next four quarters is the rational expectation implied by the policy rule (and the rest of the model),

$$i^S_{1Y,t} = 0.25E_t(i_{t+1} + i_{t+2} + i_{t+3} + i_{t+4}) \tag{3.18}$$

In sum, the model comprises equations for the Phillips curve, short-run inflation expectations, the long-run evolution of inflation expectations (imposing convergence to the central bank's inflation target), the IS curve, short-run unemployment expectations, the long-run evolution of unemployment expectations (again imposing convergence to the natural steady-state of zero for the unemployment gap), two equations defining the monetary policy rule, and a term structure equation that defines the one-year rate as the expectation of the four quarterly short-term (policy) rates. The equations in question are found in (3.3), (3.4), (3.7), (3.10), (3.12), (3.13), (3.15), (3.16), (3.17) and (3.18). As constituted, the model is a full general equilibrium model of prices, output and interest rates, and is thus suitable for counterfactual policy exercises, forecasting, and economic "story-telling," subject to the usual caveats.[19]

4. Reduced-Form and Partial Equilibrium Evidence on the Usefulness of Survey Expectations

We begin by presenting a number of single-equation results linking survey expectations measures with key macroeconomic aggregates, using multivariate relationships that are similar to those that appear in standard macroeconomic models. The point is not to claim structural

[19] Those caveats boil down to (a) incomplete confidence in the stability of all model parameters across policy regimes, and related (b) incomplete trust that we have identified truly structural relationships in all of the model's components.

identification, but to demonstrate the strong correlations between survey variables and key macro variables in regression equations that evoke standard macroeconomic relationships. We focus on the key building blocks of the simplified DSGE model laid out in section 3: A price-setting Euler equation, an "IS" curve that is motivated by a consumption Euler equation, and a monetary policy rule that is explicitly forward-looking.

Price-Setting

We estimate equation (3.3), the expectations-augmented Phillips curve. The survey expectation is the four-quarter change in total CPI inflation from the Survey of Professional Forecasters.[20] Actual inflation is measured as 400 times the log change in the total CPI, since this is the measure to which the survey expectations refer. The estimation sample is 1982:Q4 to 2014:Q4, chosen because some of the survey data are not available until the early 1980s. We employ ordinary least squares (OLS) estimation, as the survey expectations are recorded in the middle of quarter *t,* and thus can only contain price and output information for quarter *t-1* and earlier.[21]

The regression results and summary statistics are reported in panel 1.1 of table 1 below. Figure 1 displays the fitted values. Both the unemployment gap, measured as the difference between the civilian unemployment rate and the CBO's estimate of the NAIRU, and the two relative price variables for the log change in energy and food prices (dp^{e} and dp^{f}) enter contemporaneously and with two lags.[22] [23] Restricting inflation expectations to enter with a coefficient of one is a step toward

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[20] Previous studies often use this variable as a proxy for inflation expectations in such equations. The results have been replicated using the one-quarter-ahead expectation, which is available in the Survey of Professional Forecasters database and which corresponds more closely to the theory model. The results are the same in all essential respects.

[21] Replicating these equations using the one-quarter lag of expectations variables preserves the conclusions presented in the paper in all essential respects.

[22] Lag lengths are chosen using standard criteria, specifically by minimizing the AIC and Schwartz-Bayes criteria, which suggest a lag length of one or two quarters.

[23] Most of the data in this paper are real-time data—the SPF forecasts are not revised, and neither is the CPI inflation measure, the federal funds rate, or the 10-year Treasury yield. The unemployment rate has small and mostly seasonal adjustment-related revisions. The CBO's estimate of the NAIRU is not a real-time estimate. The CBO publishes quasi real-time estimates of the NAIRU back to 1991. However, from 1991 to 2009 these were updated each year only in January. Thus no real-time data are available prior to 1991, and the yearly updates create some undesirable

a more structural equation; moreover, the p-value for the F-test of this restriction is 0.76, so it is clearly not rejected by the data. The results suggest a prominent role for survey expectations in the inflation equation. These results are similar to those reported in Fuhrer (2012) and Fuhrer, Olivei, and Tootell (2012).

An empirical fact that has dogged researchers for decades is the dependence of macro variables on their own lagged values, after accounting for the normal structural influences. This empirical regularity has given rise to the inclusion of rule-of-thumb pricing or indexation (Galí and Gertler 1999; Christiano, Eichenbaum, and Evans 2005) for price-setting, and to habit formation (Fuhrer 2000; Carroll and Overland 2000) for consumption models. Table 1 shows the diminished dependence of the Phillips and IS curves on lagged dependent variables once the survey expectations are taken into account. For the estimated Phillips curve, the coefficients on the two lags of inflation are small (0.01), and are estimated imprecisely. The inclusion of additional lags further weakens statistical significance.

The autocorrelation of the residuals of equation (3.3), shown in the rightmost panel of table 1, suggests no significant autocorrelation. While this is still a somewhat reduced-form equation, it suggests little need for indexation or serially correlated markup shocks, once the survey expectations are included.

We estimate the inflation expectation equation (3.4); the results are displayed in panel 1.2 of table 1. As the table indicates, the one-year inflation expectations exhibit very strong correlation with the longer-run expectations and with the forecast for the unemployment gap one-quarter forward. The top-right panel of figure 1 shows that this simple specification captures many—but not all—of the important fluctuations in this variable over its history (note the significant residual autocorrelations for panel 1.2 in Table 1). The persistent deviations of short-run expectations from

discontinuities in the series. Replicating the key equations in table 1 suggests that the real-time series constructed using this data is dominated empirically by the current vintage of the CBO's NAIRU estimates.

the simple specification suggest sluggish adjustment of these expectations to longer-run fundamentals, a topic to which we will return in section 5.

Taken together, these two equations suggest the beginnings of a relatively coherent model for inflation. Inflation adheres to the generic form now prevalent in the literature, depending with a coefficient of one on near-term inflation expectations, and driven by a real variable whose effect is estimated with reasonable precision. Expectations in turn depend on further-out expectations of the real variable, and are anchored to the long-run inflation expectation. Thus the single-equation results point toward a structural model that can close much of the expectations loop in a way that reasonably balances theory and empirics.

IS Curve (Unemployment)

As in many studies, we find that identification of the IS curve is not straightforward. A long-term interest rate, defined as the difference between the 10-year Treasury yield and the maturity-matched inflation expectation, enters significantly in the IS equation. However, it is difficult to develop a theoretical motivation for such a relationship. The equation with the one-period short-term rate performs poorly—the estimated coefficients often have the wrong sign and quite large standard errors.

In table 1, panel 1.3a displays the OLS estimation results for equation (3.10), using the long-term real interest rate. The coefficient on the SPF unemployment gap expectation is precisely estimated at very near one, and the real interest rate is estimated with the correct sign and fairly high significance.[24] The relatively small coefficients of 0.02–0.03 on both the longer real rate and the short-term real rate are not unusual for estimated IS curves.[25]

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[24] Once again, the lag length for the real interest rate is chosen using a combination of AIC and Schwartz-Bayes criteria, which agree in this case.

[25] For many models, a 1 percentage point change in the federal funds rate is roughly equivalent to a 0.25 to 0.33percentage point rise in the 10-year rate, so the expected ratio of these coefficients is three or four to one. See, for

The R^2 for the regression is 0.99; the residual autocorrelations in the rightmost panel of table 1 suggest no serial correlation. The fitted values for the regression, shown in the bottom left panel of figure 1, suggest a very tight fit for the regression. A version of the model that includes a lagged unemployment gap, shown in the bottom row of panel of table 1.3a, develops a small (but significant) coefficient. Thus while the role for lagged unemployment now is much diminished compared to the standard specifications with habit formation (the OLS estimate of 0.19 contrasts with that of 0.7–0.9 in many published estimates of the habit parameter), the reduced-form equation rejects the hypothesis that the lagged dependent variable has no influence.

Panel 1.4 of Table 1 reports results from OLS estimates of the short-run unemployment expectation as in equation (3.11), using the same two measures of the real interest rate. As the table indicates, identification here is weak—the short-term real rate variable enters with the wrong sign.[26] Section 5 will explore system estimates of this equation, which fare much better.

Policy Rule

The results from the OLS estimation of equation (3.16) are shown in table 1, panel 1.5. All the coefficients are estimated with correct signs and high statistical significance. Note that the long-run response coefficients for expected inflation and unemployment equal the estimated values reported in table 1 premultiplied by $\frac{1}{1-a}$, which yields 2.41 and –2.25 for the responses to inflation and unemployment, respectively.

The fit of the equation, displayed in the bottom-right panel of figure 1, is quite respectable. The estimated residuals for the equation exhibit modest serial correlation, which is not surprising given the inclusion of the lagged dependent variable. While one might wish for a policy rule that

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example, Fuhrer and Olivei (2011) for a discussion of these multipliers in a discussion of the effects of the Federal Reserve's Large-Scale Asset Purchase programs.

[26] Again, note the significant residual autocorrelations in panel 1.4 of Table 1.

does not require interest-rate smoothing, sorting out the sources of apparent interest-rate smoothing is a job for another paper.[27]

Cointegration

One simple explanation for the high correlation between rational expectations and realizations of survey expectations is that the two series are integrated of order one, and thus these regressions are simply uncovering a natural cointegration between a forecast and the realizations of the variable being forecasted. In this case, the exact lead-lag timing of the regression would not matter much: The exercise would be less likely to uncover an interesting dynamic macro link and more likely to reveal the general tendency for such paired series to move together at the low frequencies.

However, conventional unit root tests suggest it is extremely unlikely that the correlation between forecasts and realizations arises from a common unit root over the sample period in question, which in this case is 1982 to the present. Thus we rule out an explanation that relies on cointegration.[28]

5. System Estimates and Identification

The preceding section provided suggestive quasi-structural evidence that survey expectations may serve as very useful proxies for expectations in dynamic macro models. The surveys enter with plausible signs and magnitudes, aid in identifying key parameters, minimize the importance of lagged

[27] A number of authors have explored possible explanations for the presence of interest-rate smoothing. One hypothesis is that for rules where the policy rate depends on current inflation and output, the lagged dependent variable proxies for the central bank's attempts to filter the information in the higher frequency data. That is, the rule with the lagged dependent variable may be rewritten as a rule without a lagged interest rate, but with geometrically declining weights on infinite lags of inflation and output, which is a crude way of filtering the signal in noisy data on the central bank's goals. This paper makes explicit use of forecasts of inflation and output that presumably filter the high-frequency noise in inflation and output but it still finds a role for the lagged policy rate, a result that suggests that there is not a compelling explanation for interest-rate smoothing.

[28] The results for the augmented Dickey-Fuller test are very strong for the inflation and unemployment gap series, rejecting the presence of a unit root with p-values of 1 percent or smaller. ADF tests are weaker for the presence of unit roots for inflation and unemployment gap expectations. However, the results for the Elliott-Rothenberg-Stock and Ng-Perron tests are extremely strong for these series.

dependent variables in explaining fluctuations, and imply a limited role for complex structural shocks.

But there are two reasons why these regressions cannot claim to provide true structural identification. First, apart from the timing of the expectations variables, the equations include some contemporaneous variables which may well be subject to simultaneity bias. Stated differently, the simultaneous causation among interest rates, output, and inflation that is latent in the data cannot be identified without simultaneously estimating the policy rule (which implies causation from output and inflation to interest rates), the IS curve (which implies causation from interest rates to output), and the Phillips curve (which implies causation from output to inflation). The goal of this section is to estimate the model of section 3, simultaneously estimating these key equations, with the aim of more confidently identifying the causal linkages among the key variables.

Second, the single equations do not explicitly solve the problem of how to close the model—that is, how to solve for future values of survey expectations, as highlighted in sections 1 and 2. In this section, we implement the compromises for closing the model with survey expectations as discussed in sections 2 and 3, absent the convenience of solving out rational expectations using conventional methods.

As suggested in section 3, initial estimates suggest that whereas short-run expectations generally track movements in the long-run inflation expectation (the central bank's inflation goal), short-run inflation and unemployment expectations appear to adjust somewhat sluggishly to longer-run expectations. More precisely, the initial estimates of equations (3.4) and (3.11) show that the one-quarter expectations persistently deviate from the fundamentals specified in the respective equations. To illustrate the issue, figure 2 shows the difference between one-quarter (SPF) and long-run inflation (SPF) and unemployment (Blue Chip) survey expectations. As the figure indicates, the gaps persist for quite a few years, also indicative of slow adjustment of short-run expectations to longer-run fundamentals.

Table 2 displays Bayesian (and, for comparison, OLS) estimates of all the parameters in the model, along with summary statistics for the simulated posterior distribution.[29] [30] Figure 3 plots the associated parameter distributions, along with prior distributions, for each of the parameters. The system estimates generally do not differ too dramatically from the OLS estimates presented in table 1. But some differences are worth noting. The key elasticities in the Phillips and IS curves, π^{u} and u^{ρ}, both increase in magnitude relative to the single-equation OLS estimates. The difference is particularly striking for the Phillips curve, which is about four times the size of the OLS estimate. Still, in both cases the standard deviation of the posterior distribution is large enough to admit either OLS or Bayesian estimates.

Critically, the estimates of intrinsic expectations persistence—the partial adjustment coefficients implied by the estimates of $(1-\mu^{\pi})$ and $(1-\mu^{U})$ in equation (3.5) and (3.12)—are large and statistically significant, at 0.86 and 0.88 respectively, with standard errors of 0.15 and 0.12. We can use the estimated model to assess the economic significance of the partial-adjustment. The top four panels of figure 5 display the in-sample fit of the model at the estimated values of $[\mu^{\pi}, \mu^{U}]$ and at values that drastically reduce the importance of partial-adjustment. It is at once obvious from this figure that the fit for inflation expectations, unemployment realizations, and unemployment expectations deteriorate quite dramatically in the absence of partial adjustment.[31] The fit of the

[29] The prior distributions, not tabulated, are standard, and appear in the graphs of the simulated posterior distributions. Standard convergence statistics suggest that all the simulated posterior distribution estimates have fully converged, both jointly and individually.

[30] Note that the shock variances are assumed to have uninformative priors, so that the maximum likelihood estimates of the variances, which are implied by the other parameter estimates, the model, and the data, will equal the posteriors.

[31] The fit is computed via a static simulation of the model over the sample indicated, taking relative food and energy prices and the long-run inflation expectation as exogenous.

Phillips curve is improved, but less significantly. Thus in this model, the intrinsic persistence in expectations is crucial for explaining fluctuations in output, inflation and their expectations.

The Bayesian estimates of the policy rule coefficients are smaller than the OLS single-equation estimates.[32] Despite this, the estimates imply reasonably aggressive responses to expected inflation and unemployment gaps. The policy rule tracks the actual funds rate quite well (not shown—the plots are very similar to those in figure 1). The estimated autocorrelations of the structural disturbances, shown in figure 4, are quite similar to those developed in the single-equation estimates of section 3, although they are generally a bit smaller and even less significant. For a model that excludes most lagged dependent variables and autocorrelated shock processes, the fit, while not an explicit estimation criterion, is quite good.

Identification is not trouble-free in this model using survey expectations. While the Phillips curve parameters are generally estimated with reasonable precision, including the parameters that govern the evolution of the inflation expectations that enter the Phillips curve, identification of the IS curve's slope is still less robust than would be ideal. For example, changing the horizon of the unemployment expectation in the equation can flip the sign of the real interest rate coefficient, which suggests that the model's ability to distinguish between the IS curve (in which one expects a negative correlation between interest rates and activity) and the policy rule (in which one expects a positive correlation) is not perfect.

Tests for the Importance of Habits and Indexation

We first run a simple single-equation omitted variable test for the exclusion of the lagged dependent variables that proxy for habits or indexation in the IS and Phillips curves. The test takes the form

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[32] Note that for comparability with the OLS estimates, the inflation coefficient reported in table 4 is the sum of i^{π} and the implicit unit coefficient on the long-run inflation expectation. The OLS coefficient is the long-run coefficient reported in section 2.

$$\varepsilon_t^i = \lambda y_{t-1}^i + \beta_i X_t^i + e_t \ , \tag{4.1}$$

where ε_t^i is the estimated structural disturbance for the inflation or unemployment gap equation, y_{t-1}^i is the lagged value of one of these variables, and the term $\beta_i X_t^i$ represents the other variables that enter the equation. The coefficient of interest is λ, and the null hypothesis is that $\hat{\lambda} = 0$, suggesting no additional role for the lagged variable. The top panel of table 3 presents the results for estimating this equation on the estimated shocks for the Phillips and IS curves. As the table indicates, consistent with the results in section 3, the coefficient on lagged inflation in the Phillips curve is estimated to be quite small and insignificantly different from zero. The coefficient on the lagged unemployment gap in the IS curve is modest and significantly different from zero, suggesting a possibly statistically important degree of habit formation.

A systems-based testing method allows joint estimation of the effect of lagged inflation and lagged unemployment in the key equations, along with the partial-adjustment mechanism for unemployment expectations represented in equation (3.12). This entails modifying the Phillips curve and IS equations as follows:

$$\begin{aligned} &\pi_t = \pi^L \pi_{t-1} + (1-\pi^L)\pi_{t+1,t}^S - \pi^u (U_t - U_t^*) + w^e dp_t^e + w^f dp_t^f \\ &U_t - U_t^* = u^L (U_{t-1} - U_{t-1}^*) + (1-u^L)(U_{t+1,t}^S - U_{t+1}^*) + + u^\rho (\rho_t - \bar{\rho}) \end{aligned} \ , \tag{4.2}$$

and estimating the parameters π^L, u^L to assess the importance of lagged dependent variables in the augmented model, pitting the lagged variables against the persistence that may be induced by the inclusion of survey expectations.

The posterior modes of the Bayesian estimates of π^L, u^L in equations (4.2) are displayed in the middle panel of table 3, along with simulated standard errors from the posterior distribution. These estimated parameter distributions suggest no role for lagged inflation but a modest role for lagged unemployment in the model. Importantly, note that the estimated partial adjustment coefficients for both inflation and unemployment expectations remain high at about 0.8, with standard errors of 0.15-0.17.

While statistically significant, how important is the lagged unemployment gap in explaining model dynamics? The bottom panel of figure 5 displays its economic significance by simulating the model at the parameter values estimated in the top panel of table 2 (the dashed black line) and alternatively by setting u^L to 0.01 (the red line). As the figure indicates, the difference in the simulated values is virtually nil. As compared to the striking impact of the partial adjustment in unemployment expectations displayed in the top panels of figure 5, the test and the simulation together suggest there is no economically significant role for lagged actual data in the model with survey expectations.

Overall, these findings are striking, and stronger than the single-equation tests. The simple OLS tests for omitted variables suggest at best a small role for lagged unemployment in the IS curve. The system tests also suggest an economically insignificant role for lagged dependent variables in the model. But the role of sluggishly-adjusting expectations in explaining current expectations appears critical to the model's success in explaining both expectations *and* realized data dynamics. This is a key finding: What had appeared to be a strong dependence on lagged endogenous variables in DSGE models is better represented as the presence of inertia in expectations. Thus the model with survey expectations is able to distinguish clearly between the direct effect of lagged realizations (as in habits and indexation) and the role of persistent expectations in determining output and inflation. In models with survey expectations, the former is far less important, while the latter takes on a key role.

A Rational Expectations "Horse Race"

The paper now examines a head-to-head comparison of DSGE models based on rational expectations with those based on survey expectations, similar to the exercises in Del Negro and Eusepi (2010) and Nunes (2010). A simple way to perform such a comparison is to augment the model equations (3.3) and (3.10) so that rational expectations enter with weight λ, and lagged dependent variables and survey expectations enter with weights as in equation (4.2), all of which sum to one as follows:

$$\begin{aligned} &\pi_t - \lambda\Pi^S_{LR,t} = \lambda(E_t\pi_{t+1} - \Pi^S_{LR,t}) + \pi^L\pi_{t-1} + (1-\lambda-\pi^L)\pi^S_{t+1,t} - \pi^u(U_t - U_t^*) + w^e dp_t^e + w^f dp_t^f \\ &U_t - U_t^* = \lambda E_t(U_{t+1} - U^*_{t+1}) + u^L(U_{t-1} - U^*_{t-1}) + (1-\lambda-u^L)(U^S_{t+1,t} - U^*_{t+1}) + u^\rho(\rho_t - \bar{\rho}) \end{aligned} \quad .(4.3)$$

Under the null hypothesis that $\lambda = 0$, the rational expectations are unimportant in the determination of the model's key variables. As λ goes to 1 and π^L and u^L go to 0, only the rational expectations matter, and the survey expectations (and equations that determine their evolution) are irrelevant. Note that in this case, the equation is structured so that long-run expectation $\Pi^S_{LR,t}$ enters the Phillips curve with a coefficient of minus one as in Cogley and Sbordone (2008).[33] As $\lambda + \pi^L$ goes to 0, the weight on the survey expectations goes to 1, and similarly for the IS curve. It is critical to note that in this parameterization, the data can choose any combination of λ, π^L and u^L, so that outcomes can include rational expectations with no lags or surveys, rational expectations with lags due to habits and indexation, survey expectations with no lags, and so on. Thus this test puts rational expectations on a completely equal footing with survey expectations. Note also that whenever rational expectations receives a non-zero weight, the rational expectations in the model take account of both the lagged dependent variables and the presence (if any) of survey expectations and their implied dynamics.

As shown in the bottom panel of Table 3, estimating this model over the sample period 1984:Q1-2007:Q3, we obtain a posterior maximum estimate for λ of 0.14, although the estimate is

[33] The precise role that trend inflation should play in a rational expectations Phillips curve remains somewhat controversial. A version of the test that excludes Π^S_{Lt} from the Phillips curve in the test equations delivers the same results: λ is estimated to be 0.11 with an equal-sized standard error, and the other parameters attain approximately the same values.

insignificantly different from zero given the standard error of 0.20.[34] The estimated impacts of lagged inflation and unemployment are 0.082 and 0.43; their standard errors are 0.13 and 0.26 respectively. The influence of lagged unemployment is diminished somewhat relative to the estimates in Table 3. But the observation on the tiny contribution of lagged unemployment in figure 5 still applies. Overall, these estimates suggest at best a very small and economically insignificant role for rational expectations and lagged dependent variables in this model, once the information in survey expectations is taken into account.

Finally, to summarize the contributions of each of these model components to aggregate dynamics, we examine the implications for the vector autocorrelation function (ACF) that arise from omitting lagged dependent variables or sluggish adjustment of survey expectations, or increasing the weight on RE.[35] Figure 6 displays the ACF for the baseline parameters (very little weight on rational expectations, as estimated in Table 3, the solid line), setting the influence of lagged dependent variables to 0 (the solid dashed line), setting partial adjustment of survey expectations to zero (the light dotted line), and increasing the influence of rational expectations to 0.95 (the asterisks). As the figure indicates, eliminating partial adjustment in survey expectations or imposing rational expectations causes significant deterioration in the ACF relative to the estimated baseline. Including lags (or not) makes little difference to the ACF. Thus the ACF provides another way to assess the key sources of dynamics in the model. Consistent with the other results, the sluggish adjustment of survey expectations is far more important than the inclusion of lagged dependent variables.

Model Forecasting Performance

[34] The parameters and standard errors are taken from the posterior density computed as described at the beginning of this section. The priors for λ, π^L, u^L are normal with mean 0.5 and standard deviation 0.2, which allows a small portion of the mass of the prior and posterior distributions to lie below zero.

[35] For linear models, the ACF provides a complete summary of the information in the data that is relevant for the likelihood (up to the scaling of the variables' variances).

We briefly compare the forecasting performance of three versions of the model: (1) The baseline version, with parameters as estimated in Table 2; (2) a version with "pure" rational expectations (no habits or indexation, no survey expectations); and (3) a version with one-half weight on RE and one-half weight on lags. As estimated, the model's steady-state deviates from the theoretical norm because the parameter $A^{\pi e}$ is allowed to deviate from one. In order to simulate the model with reasonable steady-state properties, we set $A^{\pi e}$ equal to one and re-estimate the remaining model parameters, so that the steady-state for the key variables is as expected.[36]

Table 4 displays the root-mean-squared errors for simulations of the model, both in-sample and out-of-sample, taking the prices of food and energy and the inflation goal as exogenous over the periods. The starting points are near the troughs of each of the past four recessions, and the simulations extend for 16 quarters. The baseline model dominates the other models for inflation forecast performance. The performance for the unemployment gap is almost always dominant: the pure RE model outperforms the other two for one forecast horizon. Generally, these results would suggest selection of the survey expectations model.

6. Conclusions

The development of DSGE models has made significant progress over the past 20 years. This paper examines the extent to which the use of survey expectations with a minimal set of expectations assumptions can address the problems in DSGE model with identification, with the inclusion of lagged dependent variables that stands in conflict with evidence from micro data, and with an excessive reliance on highly correlated structural shocks.

[36] Note that this parameter restriction falls outside the 95th percentile of the simulated posterior distribution for $A^{\pi e}$. A Wald test of this restriction rejects it convincingly. The steady-state values for the other variables are as expected—the unemployment gap is zero, the inflation rate attains the central bank's target, the real interest rate equals the equilibrium real interest rate, the Fisher equation holds in the long run, and so on.

The paper suggests that the improvements afforded by using surveys of forecasts as the model's expectations are substantial. First, these expectations serve well as expectations proxies in the standard linearized first-order conditions that make up DSGE models. Second, all of the results in sections 4 and 5, but particularly the tests in section 5 suggest that most all of the inertia imparted by lagged variables in previous DSGE models is better represented by inertia in expectations processes, both for inflation and for output. Thus survey-based expectations essentially eliminate the need for adding *ad hoc* model features such as indexation and habit formation, both of which have, at best, limited support in the micro data. Third and related, using survey expectations obviates the need to incorporate complex error processes into models in order to match the dynamic properties of macro data. Fourth, survey expectations perform well in a system context, allowing one to identify key parameters quite well, although it would be overly optimistic to suggest that all identification issues are solved. Fifth, in a head-to-head empirical test survey expectations strongly dominate rational expectations in a DSGE model. Finally, the paper provides methods for endogenizing survey expectations in DSGE models, relying on a minimal set of assumptions to incorporate survey expectations in a manner generally consistent with standard macro theory.

Better understanding why survey expectations respond sluggishly to fundamentals and determining whether these expectations properties are stable across policy regimes is the subject of future work. Fuhrer (2015) examines the individual responses to the SPF and Michigan surveys and finds that individual expectations are strongly anchored to the lagged central tendency of expectations, a finding that is consistent with the partial adjustment of expectations that is found in the aggregate data in this study. That work suggests that at least in an empirical sense, sluggish expectation adjustment is "micro-founded."

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Table 1 **Regression Results, Simple Single-Equation Models** **1982:Q4-2014:Q4**				
1.1 Inflation				
Variable	**Coefficient**	***p*-value**	**Residual autocorrelations**	
Expected inflation	1	(imposed)	**Lag**	**Value**
Unemployment gap	-0.12	0.018	1	0.19
Change in food prices	0.085	0.0035	2	0.088
Change in energy prices	0.10	0.00	3	0.077
Lagged inflation	**-0.021**	**0.81**		
R^2: 0.86				
1.2 Inflation expectations				
Long-run inflation expectation	0.75	0.00	1	0.74*
One-quarter-ahead unemployment expectation	-0.19	0.00	2	0.64*
R^2: 0.90			3	0.54*
1.3a Unemployment gap – long rate				
			Residual autocorrelations	
One-quarter-ahead unemployment expectation	1.03	0.00	1	0.27*
Long-term real interest rate (R10-PTR)	0.023	0.0088	2	0.085
Lagged unemployment gap	**0.17**	**0.00**	3	0.031
R^2: 0.99				
1.3b Unemployment gap – short rate				
			Residual autocorrelations	
One-quarter-ahead unemployment expectation	1.04	0.00	1	0.26*
Short-term real interest rate $(i^S_{1Y,t} - \pi^S_{1Y,t})$	0.015	0.066	2	0.072
R^2: 0.99			3	0.027
1.4 Unemployment expectations				
Long-run unemployment expectation	1	-	1	0.97*
Short-term real interest rate, *or*	-0.11	0.032	2	0.91*
Ten-year real interest rate	0.071	0.033	3	0.85*
R^2: 0.24, 0.24				
1.5 Policy rule (Funds rate deviation) 1982:Q4-2007:Q3				
Lagged funds rate	0.80	0.00	1	0.69*
SPF 4-quarter inflation expec.	0.41	0.017	2	0.55*
SPF 1-quarter unemployment expect.	-0.43	0.000	3	0.41
R^2: 0.95				
*indicates that the autocorrelation is more than two times the standard error.				

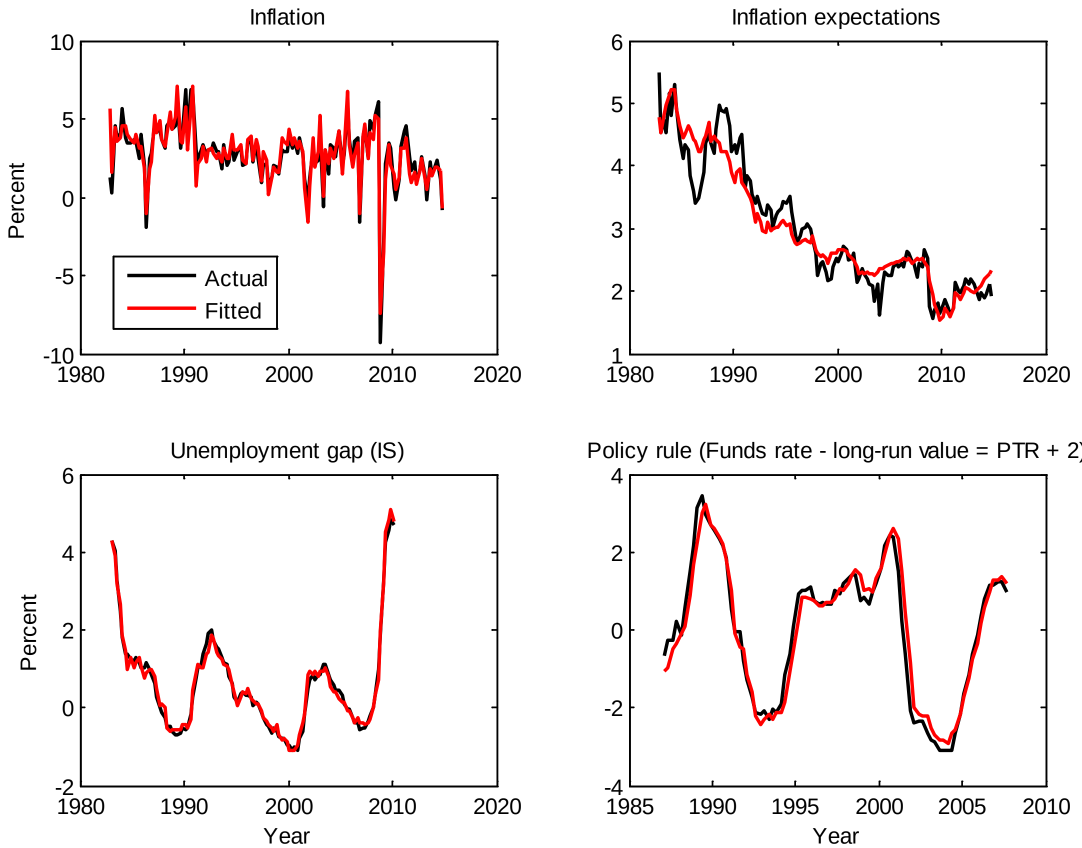
Table 2 Estimates of DSGE model with survey expectations (4 million replications)						
Posterior Distribution Summary Statistics						
Parameter	**Maximum**	**Median**	**Memo: OLS**	**Std. Dev.**	**5th percentile**	**95th percentile**
π^u	0.36	0.48	0.069	0.28	0.11	1.0
μ^π	0.14	0.20	-	0.15	0.038	0.53
$A^{\pi e}$	0.76	0.68	0.75	0.23	0.21	0.96
u^{ue}	1.1	1.1	1.03	0.16	0.74	1.3
u^ρ	0.075	0.1	0.028	0.071	0.013	0.24
$\bar{\rho}$	2.6	2.7	-	0.83	1.5	4.2
μ^U	0.12	0.18	-	0.12	0.025	0.42
a	0.83	0.79	0.81	0.14	0.52	0.97
i^π	0.97	1.1	2.4	0.54	0.27	2.1
i^u	0.41	0.76	2.3	0.49	0.14	1.7
Sample: 1984:Q2–2007:Q3 Four blocks, 1 million replications each, first 100,000 dropped for burn-in. Results for a larger burn-in allowance are virtually identical.						

Table 3 Tests of lagged variables in key macroeconomic relationships Single-equation test (equation (4.1))		
Equation	**Lag Coefficient**	***p*-value**
Phillips	0.023	0.61
IS	0.26	0.00
System Test for Importance of Lagged Variables (equation (4.2))		
Coefficient	Posterior maximum	Standard deviation
π^L	0.10	0.16
u^L	0.57	0.24
$(1-\mu^\pi)$	0.79	0.17
$(1-\mu^U)$	0.79	0.15
RE "horse race" test (equation (4.3))		
Parameter	Estimate	Standard error
λ	0.14	0.20
π^L	0.082	0.13
u^L	0.43	0.26
$(1-\mu^\pi)$	0.74	0.17
$(1-\mu^U)$	0.78	0.15

Table 4 Root-mean-squared forecast errors, various specification and samples						
	Inflation			Unemployment		
Start date	Survey	Pure RE	RE plus lags	Survey	Pure RE	RE plus lags
1984:Q2	1	3.9	1.9	0.41	1	8.2
1991:Q1	0.86	3.4	1.3	1.2	1.3	6.9
2001:Q4	1.2	2.0	1.9	0.81	0.69	4.9
2009:Q2	0.76	1.8	1.9	3.0	4.0	8.0

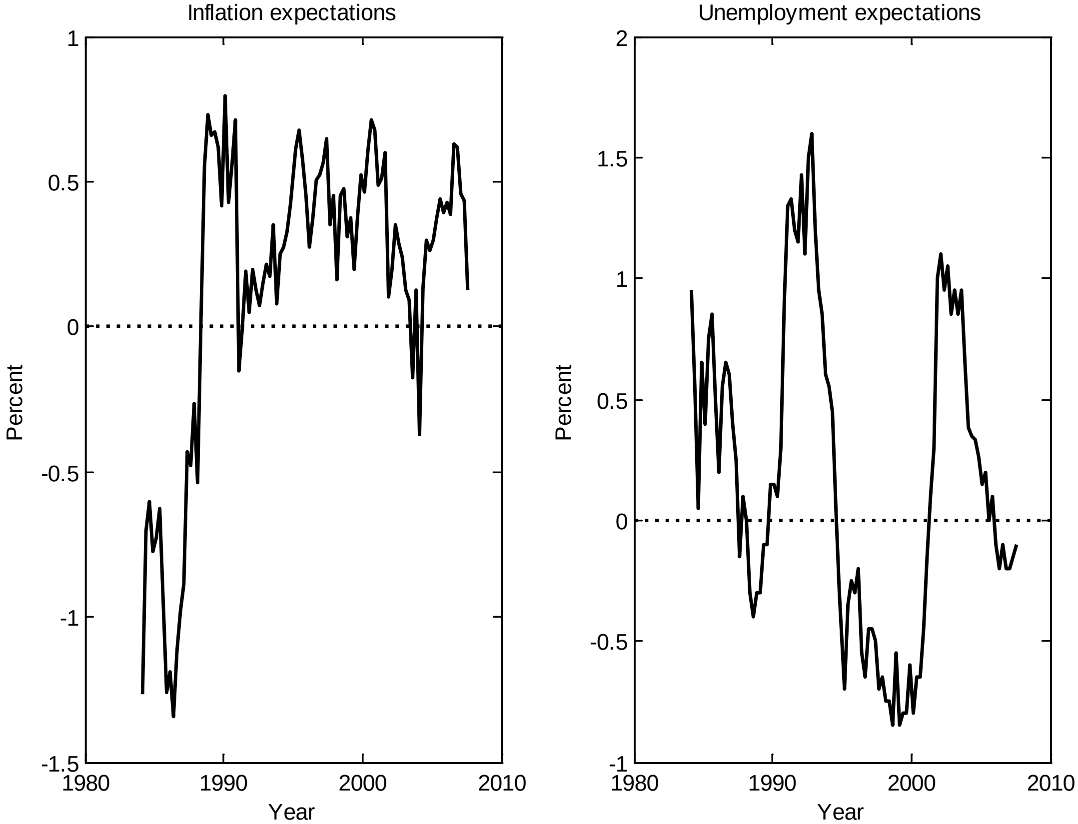
Figure 1

Fit of OLS Regressions for Key Macro Variables



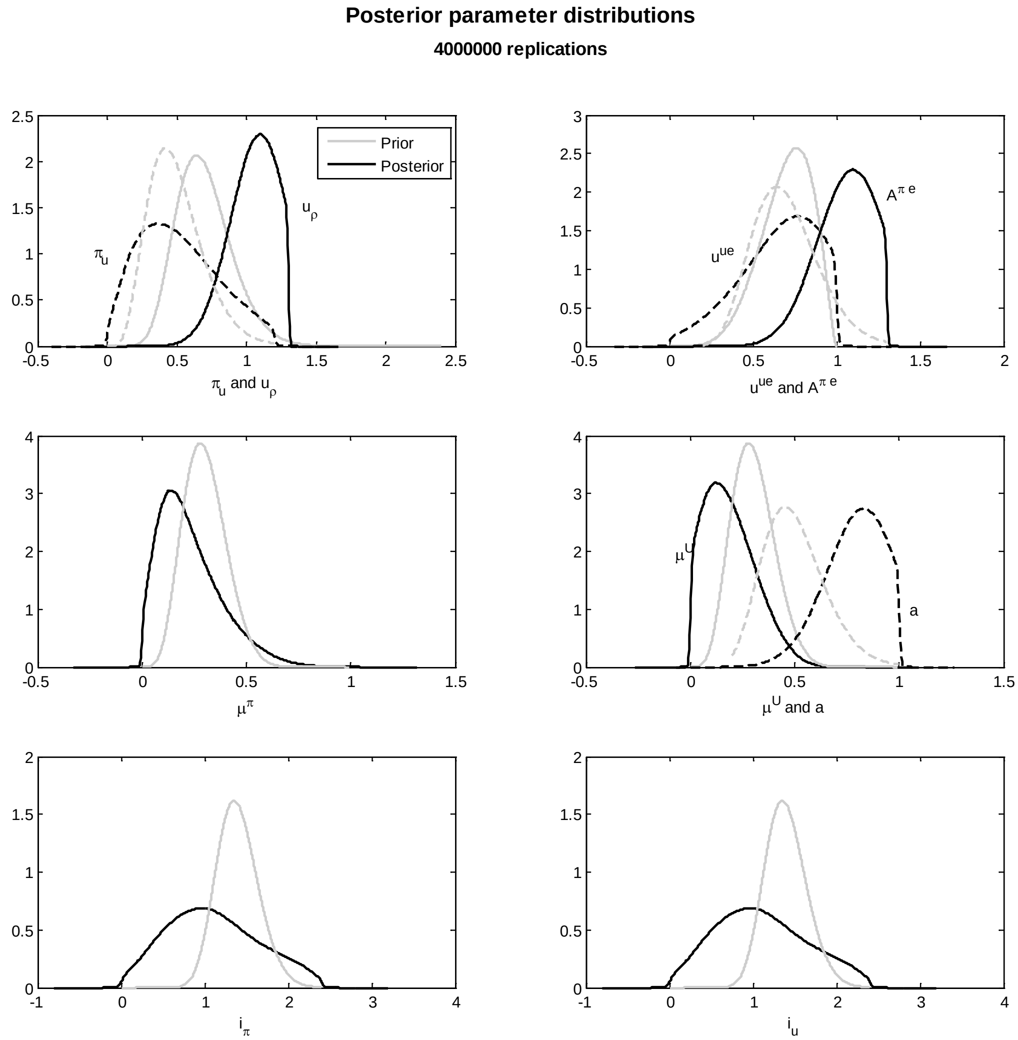
Source: Author's calculations

Figure 2
Gap Between Short-run and Long-run Expectations Measures



Source: Survey of Professional Forecasters, Blue Chip Economic Indicators

Figure 3



Source: Author's calculations

Figure 4
Autocorrelations of estimated structural disturbances

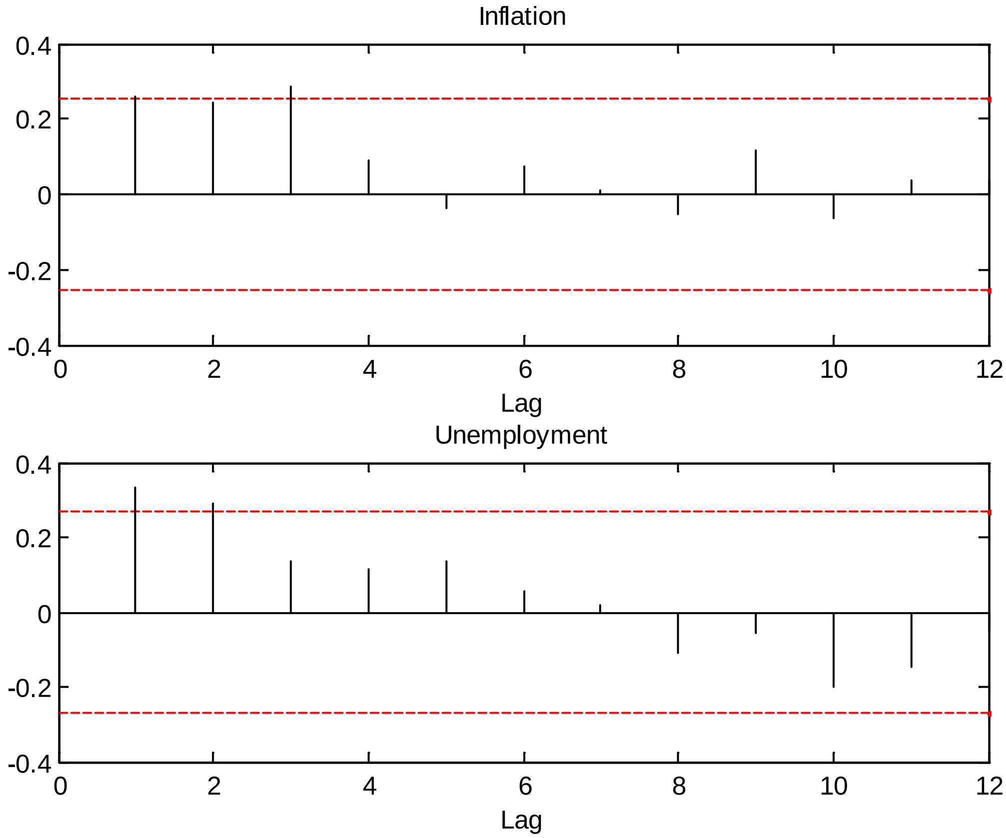
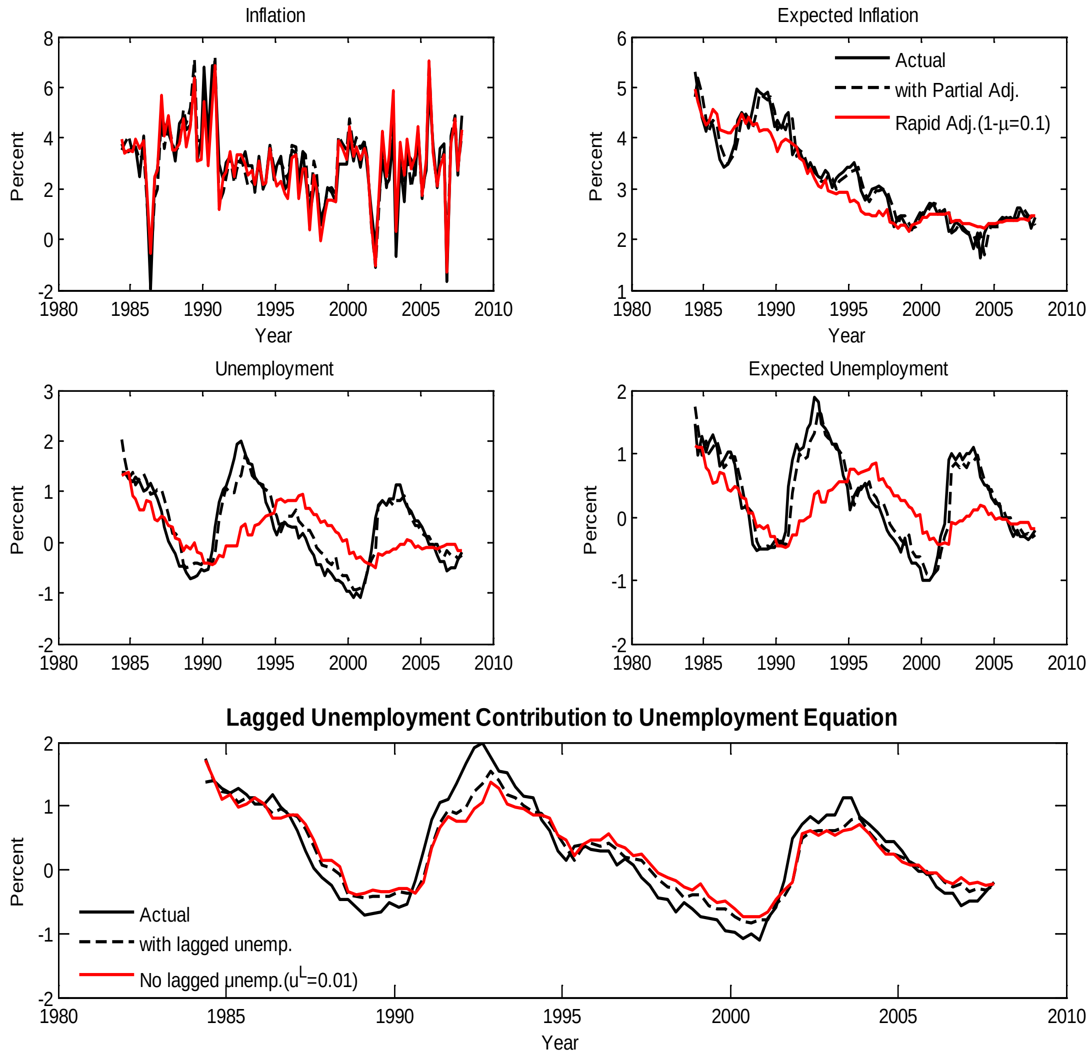


Figure 5

Importance of Partial Adjustment and Lags in Explaining Model Dynamics

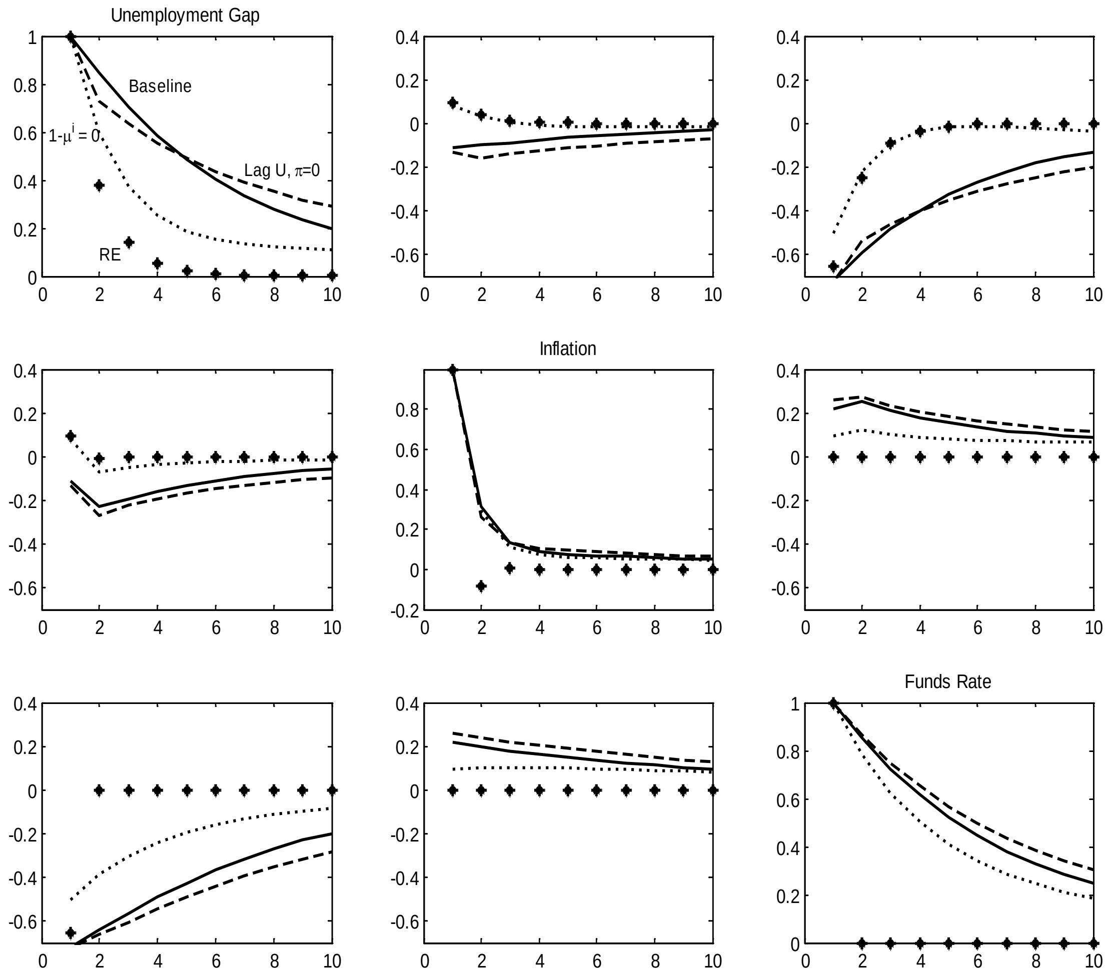
Partial Adjustment Contribution



Source: Author's calculations

Figure 6

Autocorrelation functions, with or without effects of lagged U and π, sluggish expectations or RE



Source: Author's calculations

Data Appendix

Variable	Mnemonic	Description
Inflation	π_t	400 times the log change in the total consumer price index
Unemployment rate	U_t	Civilian unemployment rate
Natural rate	U_t^*	NAIRU estimate, Congressional Budget Office (CBO)
Energy price	dp_t^e	400 times the log change in the energy sub-index of the CPI
Food price	dp_t^f	400 times the log change in the food sub-index of the CPI
Federal funds rate	i_t	Quarterly average of monthly observations of the effective federal funds rate
10-year govt. yield	R_t	Quarterly average of monthly observations of the 10-year constant-maturity Treasury yield
Survey Expectations, Survey of Professional Forecasters (SPF)		
One-year inflation expectation	$\pi_{t+1,t}^S$	Average of the median inflation expectation over next four quarters from the SPF
Long-run inflation expectation	$\Pi_{LR,t}^S$	Average of the median inflation expectation over next ten years from the SPF
One-quarter unemployment expectation	$U_{t+1,t}^S$	Median expectation of the civilian unemployment rate for the next quarter
Ten-year average unemployment expectation (less natural rate U_t^*)	$U_{LR,t}^S$	Median expectation of the average civilian unemployment rate over the next ten years, from the Blue Chip forecast survey. The March observation is taken to be the estimate of the long-run expectation for the second and third quarter of each year, while the October observation is taken to be the expectation for the fourth quarter of that year and the first quarter of the next. The variable is expressed as the deviation from the CBO estimate of the natural rate U_t^*.
One-year Treasury-bill expectations	$i_{1Y,t}^S$	Median expectation of the average 3-month Treasury bill rate over the next four quarters

Note: All SPF expectations (except the long-run unemployment expectations) are taken from the Philadelphia Fed website (http://www.phil.frb.org/research-and-data/real-time-center/survey-of-professional-forecasters/)

Observations from 1990:Q3 to the present represent information through the middle of the second month of the quarter (mid-February, mid-May, and so on.). Thus respondents will normally have one month's data on unemployment for the current quarter, no or one month's data for the CPI (depending on the CPI release date), and one month's complete data for the Treasury bill rate. Note that the dating convention used in the paper takes the month in which the surveys are returned as the current period, and all the expectations used in the paper are for the quarters following the current quarter. The timing for surveys prior to 1990:Q3 is not certain, but the Philadelphia Fed's website suggests that it is "similar."