Monetary Policy and Economic Activitiy in Japan and the United States

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Monetary Policy and Economic Activity in Japan and the United States*,**

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Abstract

A cornerstone of monetary policy making is that a looser monetary policy is associated with lower interest rates, higher growth of narrow monetary aggregates, higher output and higher inflation. These responses, which we collectively refer to as the liquidity effect hypothesis, are at odds with some of the leading theoretical models of money. This paper proposes and implements a quasi-Bayesian methodology that allows us to compare the liquidity effect hypothesis with two other hypotheses: the sticky price hypothesis and the inflation tax hypothesis. Our results indicate that there is evidence against the liquidity effect hypothesis in U.S. data, but that a skeptical Bayesian decision maker would still assign most posterior weight it. For Japan, in contrast, even a skeptic would end up favoring the sticky price hypothesis. *Journal of Economic Literature* Classification numbers: E0, E5, F0.

Key Words: Monetary Policy; Economic Activity; Japan; United States.

1 Introduction

A cornerstone of central bank policy making is that the way to stimulate the economy is to lower interest rates and thereby increase the supply of narrow money. Since Lucas's (1972) seminal article it has been generally accepted that in fact, only surprise changes in monetary policy are likely to have real effects. Still, this view of the effects of innovations in monetary policy on economic activity is so prevalent that many monetary economists evaluate the success of their models according to their ability to produce lower short term nominal interest rates, higher narrow monetary aggregates, and higher prices in response to a expansionary monetary policy shock.

In the empirical VAR literature on money, for instance, these assumptions are the starting point of identifying a shock to monetary policy[1]. When results are inconsistent with one of these assumptions, this is thought to be a shortcoming of the empirical specification. Thus when an identified expansionary monetary policy shock produces a fall in the price level it is referred to as a price puzzle.

This prevailing wisdom about the workings of monetary policy has also had a profound influence on monetary theory. For instance, the finding by Greenwood and Huffman (1987) that calibrated versions of real business cycle models with money have the property that unexpected increases in the growth rate of money increase nominal interest rates, and inflation and lower output and employment, is perceived by many to be shortcoming of this class of model (see e.g. Christiano (1991). Subsequent work by Lucas

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[1] See Bernanke and Mihov (1998), Eichenbaum, Christiano and Evans (1996) and Leeper, Sims and Zha (1996) for some recent examples.

(1990) and Fuerst (1992) was specifically motivated by a desire to overturn this counterfactual implication of flexible price models of money.

The gap between the predictions of theory and prevailing wisdom is not limited to flexible price models of money. In Rotemberg's (1996) sticky price model with monopolisitically competitive intermediate goods producers, interest rates and output both rise in response to a surprise increase in the growth rate of money.

Considerable efforts have been devoted to building theoretical models that are consistent with the prevailing wisdom. Yet success has been elusive. It is remarkably difficult to formulate either sticky price or flexible price models that produce large persistent liquidity effects without appealing to quadratic adjustment costs and/or assuming labor supply elasticities that are implausibly large (see e.g. Christiano (1991) and Christiano, Eichenbaum and Evans (1997)). Indeed, theory suggests that the liquidity effect hypothesis may not be a particularly robust phenomenon. The magnitude of liquidity effects is likely to vary across time and countries and its persistence will depend on details of the economy that we don't have much information about.

The goal of this paper is to submit this cornerstone of monetary policy to more careful scrutiny and evaluate it on an equal footing with leading alternatives that are implied by theory. Using a quasi Bayesian procedure we empirically evaluate three alternative hypotheses about the workings of monetary policy. The first hypothesis, which we will refer to as the *inflation tax hypothesis,* is consistent with flexible price cash-in-advance models of money such as Lucas and Stokey (1987), Greenwood and Huffman (1987), Cooley and Hansen (1989), and Sargent (1987). In all of these models a persistent innovation in the growth rate of money raises the nominal interest rate,

increases inflation and lowers output. The second hypothesis, which we will hereafter refer to as the *liquidity effect hypothesis*, coincides with the conventional wisdom of how monetary policy effects the economy: a surprise loosening monetary policy lowers short term interest rates, increases narrow monetary aggregates, raises output and raises the price level. The liquidity effect hypothesis is consistent with the implications of the flexible price models described in Christiano (1991) and Fuerst (1992). The third hypothesis we will consider is that an innovation in the growth rate of money acts to raise nominal interest rates, output and prices. These responses are produced by Rotemberg's (1996) sticky price model (see also Ireland (1997) and Aiyagari and Braun (1998)). At the risk of abusing the term, we will refer to these joint implications as the *sticky price hypothesis*.

We evaluate each of these hypotheses using a monte carlo based procedure proposed by Uhlig (2001)[2]. This procedure achieves identification by imposing sign restrictions directly on the impulse responses of reduced form Vector Autoregressions (VAR's) and can be given a Bayesian interpretation that allows one to calculate the posterior odds of alternative identification schemes.

Our findings indicate that there are important differences between the effects of monetary policy in Japan and the United States. In both countries we find evidence against the liquidity effect hypothesis. For the United States this evidence is not compelling. A skeptic who places high prior weights on the liquidity effect hypothesis would continue to be assign most posterior weight to the liquidity effect hypothesis after viewing our results. However, for Japan there is strong evidence that the data is more consistent with the sticky price hypothesis. This evidence would even convince a

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[2] We follow Uhlig(1999) here. But see also Faust (1999) and Canova (2001) for related approaches.

skeptical Bayesian decision maker who places a prior weight of 0.9 on the liquidity effect model and 0.05 prior weight on the sticky price model.

The remainder of the paper is organized as follows. Section 2 describes the theoretical motivation for the three hypotheses in more detail. Section 3 describes the details of our identification and evaluation procedures. Section 4 contains the results and Section 5 concludes.

2 Theoretical Motivation

This section motivates the choice of our three hypotheses regarding the effects of an innovation in monetary policy. We start by describing the inflation tax hypothesis. Monetary economists have understood that inflation acts as a tax at least since Friedman (1968). Greenwood and Huffman (1987) find that the inflation tax hypothesis is quantitatively important. They consider the dynamic effects of innovations in monetary policy in a calibrated cash-in-advance model and find that a positive innovation in the growth rate of money increases nominal interest rates, increases prices and lowers employment. In their model, inflation is a tax on labor income that induces households to work less and thus lowers output. This inflation tax effect is present in most transaction demand models of money in which there is a labor supply decision.

The second hypothesis is the liquidity effect hypothesis. While this hypothesis is the maintained hypothesis underlying most central bank actions, it is only recently that theories have been developed that produce liquidity effects in flexible price general equilibrium models. Lucas (1990) and Fuerst (1991) were some of the first researchers to

develop models that are consistent with this hypothesis. These models limit the ability of certain sectors to interact or react to an innovation in money supply. Christiano (1991) subsequently found that calibrated versions of these models often had the property that the inflation tax effect was larger than the liquidity effect. Even though a liquidity effect was present, the equilibrium responses in most cases were consistent with the inflation tax hypothesis. In addition, even when the responses were consistent with the liquidity effect hypothesis, they were not persistent and disappeared in the next period after households and firms readjusted their portfolios. Typically adjustment costs of one form or another are needed to generate persistent liquidity effects (see also Christiano and Gust (1999)).

The final hypothesis is based on the work of Rotemberg (1996). Rotemberg posits a model in which monopolistically competitive firms incur costs when they adjust their prices. A demand for money is introduced using a cash-in-advance constraint. His model successfully reproduces some of the principal empirical features of the data but has the property that a surprise increase in the growth rate of money supply raises nominal interest rates, output and prices. The reason for this is that at the time of the arrival of the shock, expectations of higher future inflation act to raise the nominal interest rate. However, prices do not fully respond to the innovation and thus current consumption is temporarily a bargain. Christiano, Eichenbaum and Evans (1997) find that this property of Rotemberg's model is robust to many natural extensions. They do succeed in producing a specification in which the nominal interest rate falls, but find that it implies a labor supply elasticity that is implausibly large and that harms the model's performance in other dimensions.

An assumption made in most of the analyses described above is that the growth rate of money supply is exogenous and persistent. This assumption is not innocuous and relaxing it could conceivably overturn all of the theoretical results documented above. Unfortunately, our understanding of how these properties of the models vary with the specification of the monetary policy feedback rule is in its infancy. However, results in Aiyagari and Braun (1998) suggest that these properties of the models may be reasonably robust to the exact details of the feedback rule. They compare and contrast simple exogenous money supply rules with optimal monetary policies in a liquidity effect model and a Rotemberg type sticky price model. In both models there is a role for an activist monetary policy. It turns out that the qualitative properties of the responses, which form the basis of our hypotheses, are the same under both the exogenous and optimal monetary policies.

Finally it should be pointed out that all of the empirical work described above is calibrated to U.S. data. We will assume below that these hypotheses are empirically relevant for Japan too.

3 The statistical model

In this section we describe the reduced form VAR's, the choice of variables and the simulation methodology used to evaluate the alternative hypotheses.

3.1 The reduced form VAR

We start from assuming the following VAR model for the macro structure:

$$x_{t+1} = C_0 + C(L)x_t + u_{t+1}, \quad u_t \sim IID(0, \Sigma) \tag{3.1}$$

where x_t is a (Kx1) vector of macroeconomic variables, L is a lag operator, and

$C(L) = C_1 + C_2 L + ... + C_J L^{J-1}$. In order to identify the innovation to monetary policy we

orthogonalize the variance-covariance matrix of u_t. That is we find a P such that

$$Px_{t+1} = PC_0 + PC(L)x_t + Pu_{t+1}, \quad E(Pu_t u_t 'P') = I. \tag{3.2}$$

The details of how P is chosen are described below. Using the transformations

$\tilde{x}_t = Px_t$ and $\varepsilon_t = Pu_t$ we can rewrite (3.2) as:

$$\tilde{x}_{t+1} = PC_0 + PC(L)P^{-1}\tilde{x}_t + \varepsilon_{t+1}. \tag{3.3}$$

3.2 Variable selection

The choice of variables for the VAR is motivated by two criteria. First, we want a list

of variables that collectively summarizes the principal links between monetary policy and

the economy. In particular, we want to include the most important variables considered

by the monetary authority when conducting monetary policy. Second, we also want the

list to include those variables that are known to produce a liquidity effect. That is we

want to bias things in favor of the conventional explanation.[3] These considerations led us

to consider two distinct lists of monthly variables for the U.S. and one list of monthly

variables for Japan. Our baseline VAR model for the U.S. consists of the six variables:

$x_t^{US} = (CPI_t, Y_t, NBR_t, R_t, TOTR_t, PCOM_t)'$ where *CPI* is the price level as measured by

the Consumer Price Index, *Y* is output as measured by Industrial Production, *NBR* is non-

borrowed reserves, *R* is the federal funds rate, and *PCOM* is a commodity price index.

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[3] Below we will report evidence against the liquidity effect hypothesis. By choosing variables that are known to be consistent with this hypothesis we are giving this hypothesis its best possible chance. This makes our evidence against this hypothesis more compelling.

We use a different baseline set of variables for Japan. First, we omit non-borrowed reserves. There is no evidence that the Bank of Japan has monitored non-borrowed reserves and in fact they do not even release data on non-borrowed reserves. In addition, efforts to construct non-borrowed reserves from existing data have the peculiar property that they are negative for substantial sub-periods of our sample.[4] Our list of variables for Japan consists of $x_t^{JP} = (CPI_t, Y_t, TOTR_t, R_t, M0_t, FX_t)'$ where, R is the call rate rate[5], $M0$ is the monetary base, and FX is the yen/\$ spot exchange rate. The yen/\$ exchange rate is included because it is an important information variable for the Bank of Japan. In order to facilitate comparison between Japan and the U.S. and to check the robustness of our conclusions for the U.S., we also report results for the U.S. using the CPI, industrial production, total reserves, the monetary base and the spot \$/yen exchange rate.

3.3 Identification of Structural Shocks

Our strategy for identifying structural shocks combines zero restrictions on the contemporaneous response of variables to structural shocks with sign restrictions on the impulse response functions.

3.3.1   Zero restrictions

We impose a block recursive structure that nests the recursive identification scheme advocated by Christiano, Eichenbaum and Evans (1998) as a special case. We partition the vector of variables into three blocks. For the U.S. baseline case, the first block consists of the price level and industrial production, the second bock includes non-borrowed reserves and the federal funds rate. The third block includes total reserves and the commodity price index. To set notation suppose that P^{-1} is block triangular:

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[4] See Shioji (2000) for more details.
[5] We use the monthly average of the overnight rate on uncollateralized loans.

$$P^{-1} = \begin{pmatrix} P_{11}^{-1} & 0 & 0 \\ P_{21}^{-1} & P_{22}^{-1} & 0 \\ P_{31}^{-1} & P_{32}^{-1} & P_{33}^{-1} \end{pmatrix} \tag{3.4}$$

All the sub-blocks of P^{-1} are dimensioned (*2x2*). Observe next that $\Sigma = P^{-1}P^{-1}{}'$

implies that Σ will have the same number and shape of partitions as P^{-1}.

The block recursive structure is reflected by the fact that the partitions above the

diagonal are all matrices of zeros. This structure imposes restrictions on the

contemporaneous responses of variables in sector *j* to shocks in sector *i*. Under these

assumptions all variables in the second and third blocks respond contemporaneously to

shocks in the price level and industrial production[6]. Shocks to non-borrowed reserves and

the federal funds rate will only have contemporaneous affects on variables in blocks two

and three. Total reserves and the commodity price index have no contemporaneous affect

on variables in the other two sectors.

Under these assumptions the task of identifying the five structural shocks comes

down to determining the sub-matrix P_{22}^{-1}. Our recursive restriction on the first block is

sufficient to pin down P_{21}^{-1} and P_{31}^{-1}. Given a particular choice of P_{22}^{-1}, P_{32}^{-1} is determined

uniquely from Σ.

The block recursive structure does impose some restrictions on P_{22}^{-1}. The elements of

P_{22}^{-1} must be chosen so that:

1) shocks to non-borrowed reserves and the federal funds rate are orthogonal.

and

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[6] Formally we can identify monetary policy without imposing any other restrictions on the (1,1) block of *P*. However, identification of monetary policy also depends on the other auxiliary assumptions relating to the block triangular structure of *P*. In Braun and Shioji (2001) we attempt to completely identify all of the shocks.

2) $P_{22}^{-1} P_{22}^{-1}{}' = \Sigma_{22} - \Sigma_{21} \Sigma_{11}^{-1} \Sigma_{21}{}' \equiv \Omega$

We will show below that these restrictions only identify P_{22}^{-1} up to a scalar.

3.2.2 Sign Restrictions

The system described above is not completely identified. In order to complete the identification of monetary policy, we impose sign restrictions on the impulse response functions. Our methodology for doing this is a rejection based quasi-bayesian monte-carlo procedure that builds on previous work by Canova (2002), Faust (1999), and Uhlig (2001).

Before going into the details, it is helpful to the reader to provide an overview of how this procedure works. We start with a set of sign restrictions on the impulse response functions that embodies one of our three hypotheses regarding the effects of monetary policy shocks on economic activity. The exact form of the restrictions and their motivation are described in section 4 below. We then randomly draw from the posterior distributions of the matrix of reduced form VAR coefficients, the variance covariance matrix of the error term, Σ, and the free elements of P_{22}^{-1} to find a set of coefficients that satisfy the sign restrictions. If a particular monte-carlo draw satisfies the sign restrictions we tabulate it, otherwise it is discarded.

Let $\hat{C}_0$, $\hat{C}(L)$ and $\hat{\Sigma}$ denote the estimated values of the coefficients and variance covariance matrix of the estimated reduced form VAR. Under a diffuse prior the coefficients' posterior will be normally distributed and the variance covariance matrix will be Wishart distributed (see Uhlig (2001) for more details). The first step is to take a draw from the posterior distribution of coefficients and variance covariance matrix of the VAR. Denote the ith random draw by $\{\hat{C}_{0,i}, \hat{C}_i(L), \hat{\Sigma}_i\}$. A draw from the posterior

distribution of the variance covariance matrix gives us a random realization for the sub-matrix $\hat{\Sigma}_{22,i}$ and a realization of $\hat{\Omega}_i$ given in (3.4). Next, we calculate the eigenvalues and eigenvectors of $\hat{\Sigma}_{22,i}$ and $\hat{\Omega}_i$ and perform a second monte carlo simulation over the free elements in P_{22}^{-1}.

Take $\hat{\Sigma}_{22,i}$ and denote the eigenvalues of this $(2x2)$ matrix as μ_1 and μ_2, and the corresponding eigenvectors as v_1 and v_2. Uhlig(2001) shows that the first column of P_{22}^{-1}, which we denote by a, has to take the following form:

$$a = \sum_{m=1}^{2} \alpha_m \cdot \sqrt{\mu_m} \cdot v_m \tag{3.5}$$

where the α's are weights attached to each of the two eigenvalues. We impose the following normalization:

$$\sum_{m=1}^{2} \alpha_m^2 = 1. \tag{3.6}$$

This leaves us with one degree of freedom to determine the weights. We draw α_1's randomly from a uniform distribution, and then choose α_2's to satisfy condition (3.6). An α chosen in this way pins down the first column of P_{22}^{-1}. The second column is calculated using the restriction $P_{22}^{-1} P_{22}^{-1}{}' = \Sigma_{22} - \Sigma_{21}\Sigma_{11}^{-1}\Sigma_{21}' \equiv \Omega$. At this point we have a completely specified data generating mechanism and can calculate impulse response functions and ascertain whether or not they satisfy our sign restrictions.

We turn now to describe how sign restrictions on the impulse response functions are imposed and used to discriminate among the three hypotheses.

3.4 Imposing the three hypotheses on the data

Table 1 summarizes the sign restrictions that the three hypotheses imply for the responses of prices, output, narrow money and the interest rate following a contractionary monetary policy shock. Observe that the three hypotheses impose distinct restrictions on the impulse response functions.

Table 1

Response of macroeconomic variables to a contractionary monetary policy shock under the three alternative hypotheses[7].

	Prices	Output	Money	Interest rate
Inflation Tax hypothesis	Up	Down	Up	Up
Liquidity effect hypothesis	Down	Down	Down	Up
Sticky Price hypothesis	Down	Down	Down	Down

In order to complete the specification of the sign restrictions it is necessary to specify the horizon over which these restrictions are binding. Friedman (1968) suggested that the liquidity effect might be operative at horizons of up to a year. We choose to only restrict the responses in the first five to six months after the arrival of the shock and do so in a rather weak way. Let month 0 denote the month in which the shock to monetary policy arrives, month 1 denote the first month after the arrival of the shock and etc. For output and prices we will assume that the sign restriction for a particular hypothesis is satisfied if the impulse response function for the respective variable has the correct sign in a majority of months 1 through 6. For money and the interest rate we will assume that the

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[7] Note that we are defining a contractionary monetary policy here to be one that produces a fall in output.

hypothesis is satisfied if the impulse response function for the respective variable has the correct sign in a majority of steps 0 through 5. This distinction between prices and output, on the one hand, and money and interest rates, on the other hand, arises because the block recursive structure implies that the response of output and prices in month 0 is zero.

In choosing this particular set of sign restrictions we tried to strike a balance between two issues. First, in existing monetary models most variables respond quickly to innovations in monetary policy and responses peak within one or two months of the arrival of the shock. While these models may be lacking in propagation, they reflect our best understanding of how the economy works and we think these restrictions should be taken seriously and imposed on the data. On the other hand, the empirical VAR literature on identifying monetary policy shocks often finds that it can take up to two years for some variables, such as prices, to show a statistically significant response. To accommodate these findings, we chose to make the restrictions relatively weak and only require that a majority of the signs be correct in the first 5 to 6 months after the arrival of the shock.

Finally, it is important to note that these sign restrictions are joint restrictions on the coefficients of the VAR, the variance covariance matrix of the disturbances and the $\alpha's$. A valid data-generating mechanism consists of a draw from the posterior distribution of the estimated coefficients, a draw from the posterior distribution of the variance covariance matrix, Σ, and a particular vector of $\alpha's$ that satisfy all of the sign restrictions for a particular hypothesis. Thus, the frequency of valid draws for a particular hypothesis can also be interpreted as a measure of the model's fit.

Up to this point the discussion has allowed for uncertainty in the parameters of a

particular model, but has not allowed for structural uncertainty. There are two different types of structural uncertainty that we would like to consider. First, we would like to allow for uncertainty about which of the three hypotheses is best. Second, we would also like to allow for uncertainty about the number of lags in the VAR.

To assess the plausibility of the alternative structures, $\{S_i, i=1,2,\ldots,I\}$, we calculate the posterior probabilities of each structure given the data X using Bayes formula:

$$p(S_i \mid X) \doteq cp(S_i)p(X \mid S_i) \tag{3.5}$$

where c is a normalizing constant that insures that the probabilities sum to one, $p(S_i)$ is the prior probability of each structure, and $p(X \mid S_i)$ is the probability of the data given S_i. The frequency of successful draws for each hypothesis, which we have described how to calculate above, provides us with $p(X \mid \hat{C}, \hat{\Sigma}, S_i)$. Then using the following formula proposed by Draper (1995) we can approximate $p(X \mid S_i)$ by:

$$\ln p(X \mid S_i) = \frac{1}{2}k_i \ln(2\pi) - \frac{1}{2}k_i \ln n + \ln p(X \mid \hat{C}, \hat{\Sigma}, S_i) \tag{3.6}$$

where k_i is the number of parameters in the ith structure and n is the number of observations. This formula is particularly useful when comparing across specifications with different lag lengths and thus numbers of parameters.

4 Results

4.1 U.S. data

Table 2 reports the number of successful draws under each of the three hypotheses under four different scenarios. All results are based on a sample period running from

1981:1 through 1999:12. The total number of draws in each case was 50,000. Results are reported for VAR's with 12 lags.

Table 2

U.S. data VAR(12): Number of valid draws out of 50,000 total draws.

	Baseline: CPI, IP, NBR, R, TOTR, PCOM		Alternative: CPI, IP, TOTR, R, M0, $/yen	
	All constraints imposed	Price responses not constrained	All constraints imposed	Price responses not constrained
Liquidity Effect Hypothesis	4,812	12,047	683	1,248
Sticky Price Hypothesis	4,819	4,935	5,136	5,196
Inflation Tax Hypothesis	9,232	9,641	14,806	14,968

The first column reports results for the baseline set of variables with all of the constraints described in section 3.4 imposed. If one starts from a uniform prior over the three alternative hypotheses, the posterior odds ratios indicate that the inflation tax hypothesis is 1.9 times as likely as the other two hypotheses. The results also imply that the sticky price hypothesis and the liquidity effect hypothesis are equally plausible.

Given the strong priors that the profession has in favor of the liquidity effect hypothesis, it is interesting to ask how this empirical evidence might affect the beliefs of a Bayesian decision maker whose prior is heavily weighted in favor of the liquidity effect hypothesis. Suppose that one starts with prior beliefs of (0.9, 0.05, 0.05) over

respectively the liquidity effect hypothesis, the sticky price hypothesis and the inflation tax hypothesis, then these results imply posterior probabilities of (0.86, 0.048, 0.092). From this we see that a Bayesian decision maker would continue to be very confident in the liquidity effect hypothesis after viewing the evidence presented in column 1.

The impulse response functions and one standard error confidence intervals for the baseline results reported in column 1 can be found in Figure 1. Looking first at the results for the liquidity effect hypothesis in the first column we see that the results are broadly consistent with findings elsewhere in the literature (see e.g. Christiano, Eichenbaum, and Evans (1998) for a survey). The price response is small in early periods and then declines thereafter. Non-borrowed reserves fall sharply in early periods but damp quickly. By period 6, the response of non-borrowed reserves is insignificantly different from zero. The response of the Federal funds rate is also strongest in early periods but transient. The response of total reserves is persistently negative. And commodity prices cycle down, up and down. The response of output though is different from the previous literature. Even though we restrict the output response to be negative in a majority of the first 5 periods, output rises in the first two periods following the shock.

The results for the sticky price hypothesis and the inflation tax hypothesis are reported respectively in columns two and three. Notice that the results for these two hypotheses are quite similar with the exception of output. Output falls in early periods for the sticky price hypothesis and rises in all periods for the inflation tax hypothesis. The similarity of the responses for the two hypotheses is broadly consistent with what theory would predict. As prices adjust under the sticky price hypothesis, one would expect that

the inflation tax effect would dominate and that the responses at medium horizons would be very similar under the two hypotheses.

Finally, note that there is a substantial difference in the output response between the liquidity effect hypothesis, on the one hand, and the sticky price and inflation tax hypotheses on the other hand. Under the liquidity effect hypothesis the response of output is about zero from month 10 and on. Under the other two hypotheses the response of output is larger and more persistent. This finding is also confirmed by variance decompositions. Under the liquidity effect hypothesis monetary policy explains less than 3.5% of the variance in output at all horizons of 24 months or less. Under the inflation tax hypothesis, on the other hand, monetary policy explains 11% of the variance in output at the 12 month horizon and 15% of the variance in output at the 24 month horizon.

Do commodity prices resolve the price puzzle?

It has been argued that for U.S. data including commodity prices in the list of variables resolves the price puzzle. Our framework provides a way to assess this question. If the commodity price index resolves the price puzzle then relaxing the restriction on prices should not have much affect on the number of accepted draws or on the nature of the response of prices under the liquidity effect hypothesis. Results reported in column 2 suggest that commodity prices are not a complete resolution to the price puzzle. Notice that the number of successes for the liquidity effect hypothesis rises substantially. This shows that our sign constraint that prices have the correct sign in a majority of the first 6 periods binds even when commodity prices are included in the VAR. Plots of the average impulse response functions for this scenario are presented Figure 2. If we compare the responses for the liquidity effect hypothesis reported in

Figure 2 with those in Figure 1, they don't look all that different. Now the response of the CPI rises slightly in early periods before falling from month 7 and on. Both plots of prices are consistent with the notion that prices don't respond much in early periods following a shock to monetary policy.

4.1.1 Robustness

Much of the previous VAR literature has conditioned on the assumption that the liquidity effect hypothesis is correct. The results presented so far show some evidence against this hypothesis using monthly U.S. data. Still, a skeptical Bayesian decision maker who is reasonably firm in the belief that the liquidity effect hypothesis is correct would still assign most weight to the liquidity effect hypothesis after being presented with empirical evidence on the other two hypotheses. However, the analysis, so far, has used the same variables and the same number of lags as the previous literature. Presumably, both the choice of variables and the number of lags in the previous literature has been to some extent driven by a desire to produce impulse responses that are consistent with the predictions of the liquidity effect hypothesis.

Columns 3 and 4 of Table 2 provide some evidence on the extent to which the results depend on the particular choice of variables. These results are based on runs with 12 lags, a sample period of 1981:1 through 1999:12 and a vector of variables that consists of the consumer price index, industrial production, total reserves, monetary base and the $/yen exchange rate. We chose these variables because they include some of the more important determinants of demand and supply for reserves and base money and also other potentially important information variables e.g. the $/yen exchange rate. We have left the commodity price level out because as we noted above it doesn't resolve the price puzzle.

In addition this same set of variables is also available for Japan. This facilitates comparison of the results across the two countries. The results using this alternate set of variables offers stronger evidence in favor of the inflation tax hypothesis. Now if one starts with a prior of (0.9, 0.05, 0.05) the posterior distribution over the three hypotheses is: (0.38, 0.16, 0.46) with all constraints imposed. If the restrictions are relaxed on the response of prices the posterior distribution of the three hypotheses is: (0.53,0.12,0.35). From this we see that the results favoring the liquidity effect hypothesis are not robust to the choice of dataset. As in the baseline specification relaxing the constraints on the response of prices does not produce a price puzzle.

Table 3

U.S. Baseline specification: VAR (3), VAR(6) and VAR(12), all constraints imposed.

	3 lags	6 lags	12 lags
Liquidity Effect Hypothesis	2005	2266	4812
Sticky Price Hypothesis	15959	8626	4819
Inflation Tax Hypothesis	1284	6416	9232

Finally, to investigate the robustness to the number of lags we also repeated the simulations for the baseline model using 3 and 6 lags. The results for the baseline parameterization are reported in Table 3. An inspection of this table indicates that the

results are not robust to the number of lags. For both the three lag and six lag specifications the sticky price hypothesis has the highest number of valid draws. Using formula (3.6) we can also do comparisons across alternative lag lengths. Applying formula (3.6) to calculate $p(X \mid S_i)$ effectively rules out *all* of the specifications with 6 or 12 lags. This is because the large number of additional parameters in the 6 and 12 lag specifications reduces their conditional probabilities by several orders of magnitude. According to this metric for correcting for degrees of freedom the only results for the baseline specification that are relevant are those reported in column 1 of Table 3. If we start with a prior of (0.9, 0.05, 0.05), the posterior odds are respectively (0.68, 0.3, 0.02). Thus, a skeptical Bayesian decision maker would still assign posterior odds of about 2 to 1 in favor of the liquidity effect hypothesis[8].

Overall, the results presented here indicate that if one conditions on the variables generally used in the empirical VAR literature and assigns high prior probability to the liquidity effect hypothesis, it is not possible to rule it out as the most likely of the three hypotheses for the United States.

4.2   Results for Japan

For Japan we consider a VAR that includes the CPI less food, industrial production, total reserves, the call rate, M-0 and the yen/$ exchange rate. The sample period starts in 1981:1 and ends in 1996:12. We chose to end the sample period here because there were several unusual events that occurred in 1998-1999. In 1998 markets for overnight interest rates were disrupted due to concerns about default by Japanese banks. The Bank of Japan's zero interest rate policy caused further disruptions in 1999. Results for Japan

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[8] We also simulated 3 and 6 lag specifications in which prices were not restricted. However, these specifications produced persistent price puzzles that lasted 17-24 months.

using 3, 6 and 12 lags in the VAR are reported in Table 4.

Table 4

Japan: Number of valid draws out of 50,000 total draws for alternative lag lengths and alternative identifying assumptions

	All constraints imposed			No constraints on response of prices		
	3 lags	6 lags	12 lags	3 lags	6 lags	12 lags
Liquidity Effect Hypothesis	1	237	1,138	21,087	15,845	12,220
Sticky Price Hypothesis	6,540	14,988	23,802	23,454	24,151	28,309
Inflation tax Hypothesis	2,664	3,429	1,770	2,985	3,651	1,801

Consider first the results with 12 lags reported in column 3. Notice that the sticky price hypothesis has the highest number of successful draws. Even if one starts out with priors of (0.9, .05, .05), the posterior odds of (0.44, 0.52, 0.04) favor the sticky price hypothesis. The impulse responses for the specification with 12 lags are reported in Figure 4. A comparison of the second and third columns in Figure 3 reveals an important difference between the sticky price and inflation tax hypothesis results. The output responses under the two hypotheses are quite different. Under the sticky price hypothesis the response of output is negative for 20 months before turning positive. Under the inflation tax hypothesis, in contrast, the response is positive in all months except month 2.

There are also some differences in the response of exchange rates across the three hypotheses. Under the liquidity effect hypothesis the response of the exchange rate is generally negative indicating nominal appreciation of the yen while under the other two hypotheses the yen appreciates in the impact period and then depreciates in all subsequent periods. However, it is difficult to assess whether these exchange rate

responses are consistent with theory without imposing other restrictions from theory such as uncovered interest rate parity[9].

Returning to Table 4, consider next columns one and two that report results with three and six lags. Here the evidence against the liquidity effect hypothesis is even more compelling. As was the case for U.S. data, adjustments for degrees of freedom using equation (3.6) rule out all specifications with six and twelve lags and lead one to assign all the mass of the posterior distribution to the results with three lags. For the three lag specification the posterior probabilities associated with a prior of (0.9, 0.05, 0.05) are (0, 0.71, 0.29). The liquidity effect hypothesis has a posterior probability of about zero. These results indicate that the sticky price hypothesis is the best description of how monetary policy affects the macroeconomy in Japan.

Consider next column 6, which reports results for a 12 lag specification with no restrictions imposed on the response of prices. Here the performance of the liquidity effect hypothesis is substantially better. Relaxing this restriction on prices increases the number of successful draws from 1,138 to 12,220. However, this is not without a cost as can be seen in Figure 5. Relaxing this restriction also produces a persistent price puzzle. Prices have the wrong sign in each and every one of the first 24 months following the arrival of the shock. Persistent price puzzles also occur in the unrestricted 3 and 6 lag specifications. Inasmuch as a persistent positive price response is inconsistent with the implications of the liquidity effect hypothesis, these results are inadmissible.

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[9] We were unwilling to do this given the considerable evidence against uncovered interest rate parity. (See e.g. Engel (1995) for a survey of the literature on this topic.)

4.2.1 Robustness

In results not reported here we have also experimented with other lists of variables. One case of particular interest is whether including commodity prices resolves the price puzzle for Japan. This does not occur. We performed a simulation for a 12 lag specification in which the variables were the Consumer Price Index less food, industrial production, total reserves, the call rate, M0 and a commodity price index. For the liquidity effect hypothesis the number of successful draws fell from 1,138, as reported in column 3 of Table 4, to 794. The Japanese price puzzle is not resolved by including commodity prices.

McCallum (1994) has argued that the spread on long and short rates is an important information variable for the monetary authority so we also re-estimated the three models using CPI less food, industrial production, M0, the call rate rate, the yield on 10 year offshore swaps[10] and the yen/dollar exchange rate. For a specification with 12 lags the posterior distribution associated with our skeptical prior was (.04, .80, .6). We also extended the sample to 1999:12 and found that including this additional data had no substantial effect on the results. Finally, we also tried runs using the 1 month Tibor rate instead of the call rate and found that this also had no substantive effect on our results.

Taken together these results indicate that Japanese data is inconsistent with the liquidity effect hypothesis and most consistent with the sticky price hypothesis.

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[10] Open market purchases and sales of long-term bonds are large in Japan- about 70% of monetary base in Japan is backed by long-term bonds.

5 Concluding Remarks

A cornerstone of central bank policy in most countries is that an expansionary monetary policy is associated with the liquidity effect hypothesis. Results presented here suggest that this premise should be viewed with caution. Using monthly data we have found that it is difficult to dismiss this hypothesis using monthly U.S. data. However, this empirical support is not robust to changes in the list of variables.

For Japan, it is much more difficult to reconcile monthly data with the liquidity effect hypothesis. The sticky price hypothesis performs better across a wide variety of specifications.

It's worth noting that our results are not of necessarily at odds with the liquidity effects in high frequency Japanese data on bank reserves.[11] Hayashi (2000), for instance, has found empirical evidence of liquidity effects at the end of reserve maintenance periods. If periodic unexpected shocks to bank's reserves occur towards the end of the maintenance period, this can induce a precautionary demand for liquidity. However, these effects disappear at the start of the next maintenance period because banks reserve requirements are based on average balances over the entire maintenance period and they thus have great flexibility in adjusting their reserve balances early in the maintenance period. Embedding these types institutional details of the Japanese market for reserves in a general equilibrium hypothesis that can link these types of liquidity effects to movements in larger monetary aggregates or other macro variables is an interesting topic for future research.

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[11] Note that this use of the term liquidity effect is more narrowly defined as an increase in the supply of reserves that drives the overnight rate down.

More generally we view the empirical methodology we have described here to be an attractive way to incorporate identifying restrictions from theory. In related work we condition on the sticky price hypothesis for Japan and investigate the response of the Japanese Term Structure of Interest rates to innovations in monetary policy. We are also exploring generalizations of Uhlig's identification scheme in which we attempt a more general identification of macro shocks in an asset-pricing framework.

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Figure 1

U.S. baseline responses for liquidity effect, sticky price, and inflation tax models.

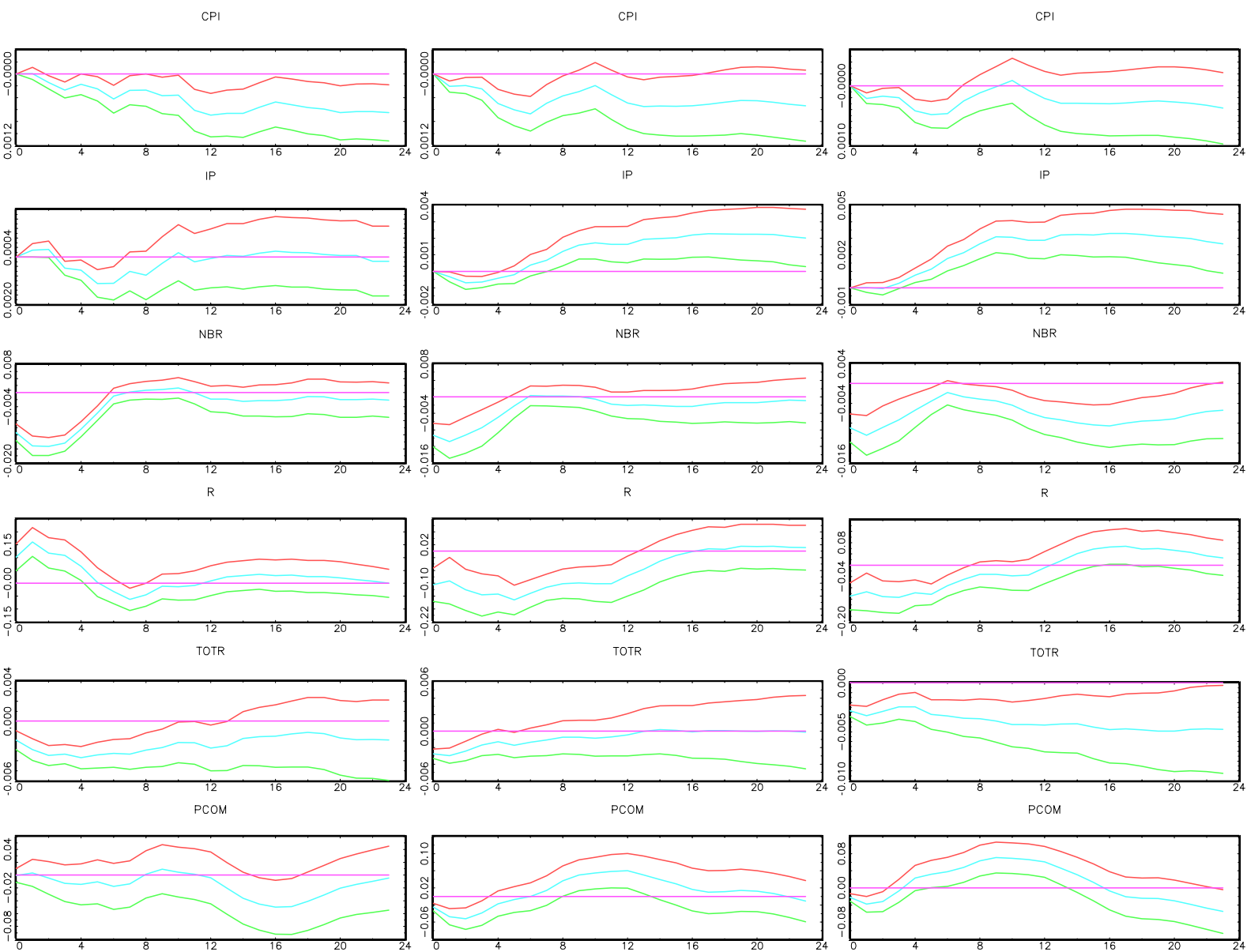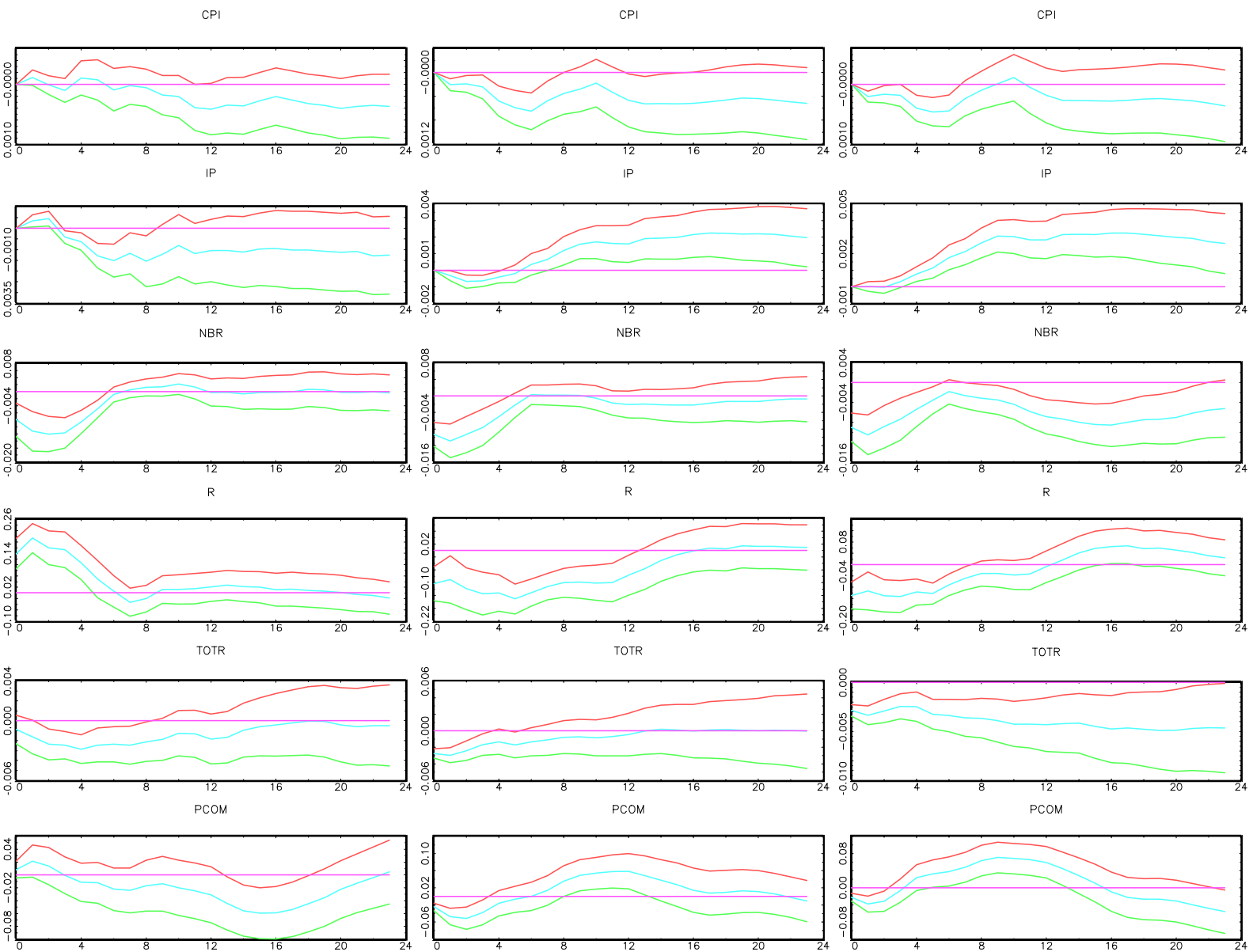| Liquidity Effect Model	Sticky Price Model	Inflation Tax Model

Figure 2

U.S. responses for liquidity effect, sticky price and inflation tax models. Price responses not constrained.

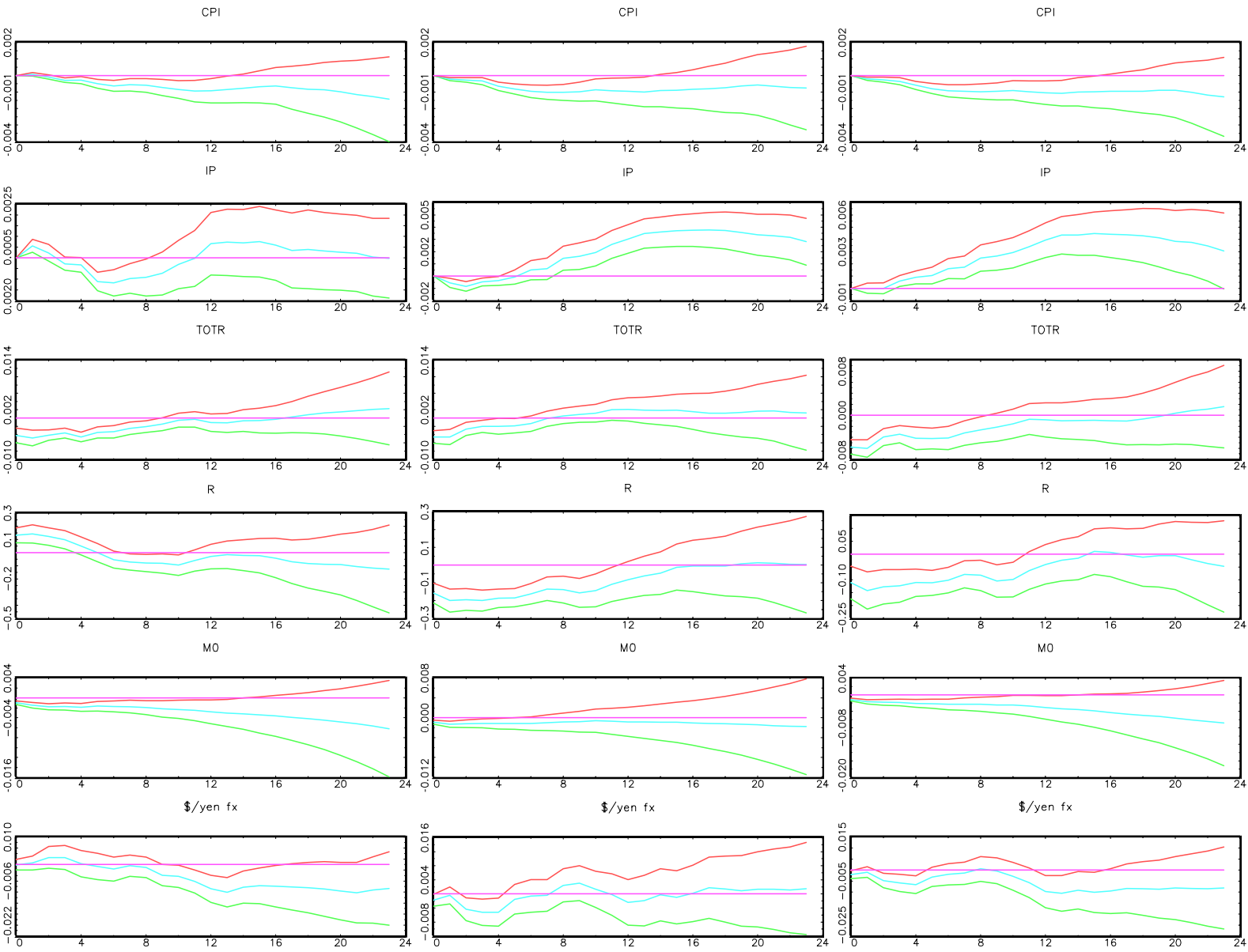Liquidity Effect Model

Sticky Price Model

Inflation Tax Model

Figure 3
U.S. Alternative specification responses for the three models.

Liquidity Effect Model · Sticky Price Model · Inflation Tax Model

CPI · IP · TOTR · R · M0 · $/yen fx

Figure 4
Japan Baseline responses of three models

Liquidity Effect Model
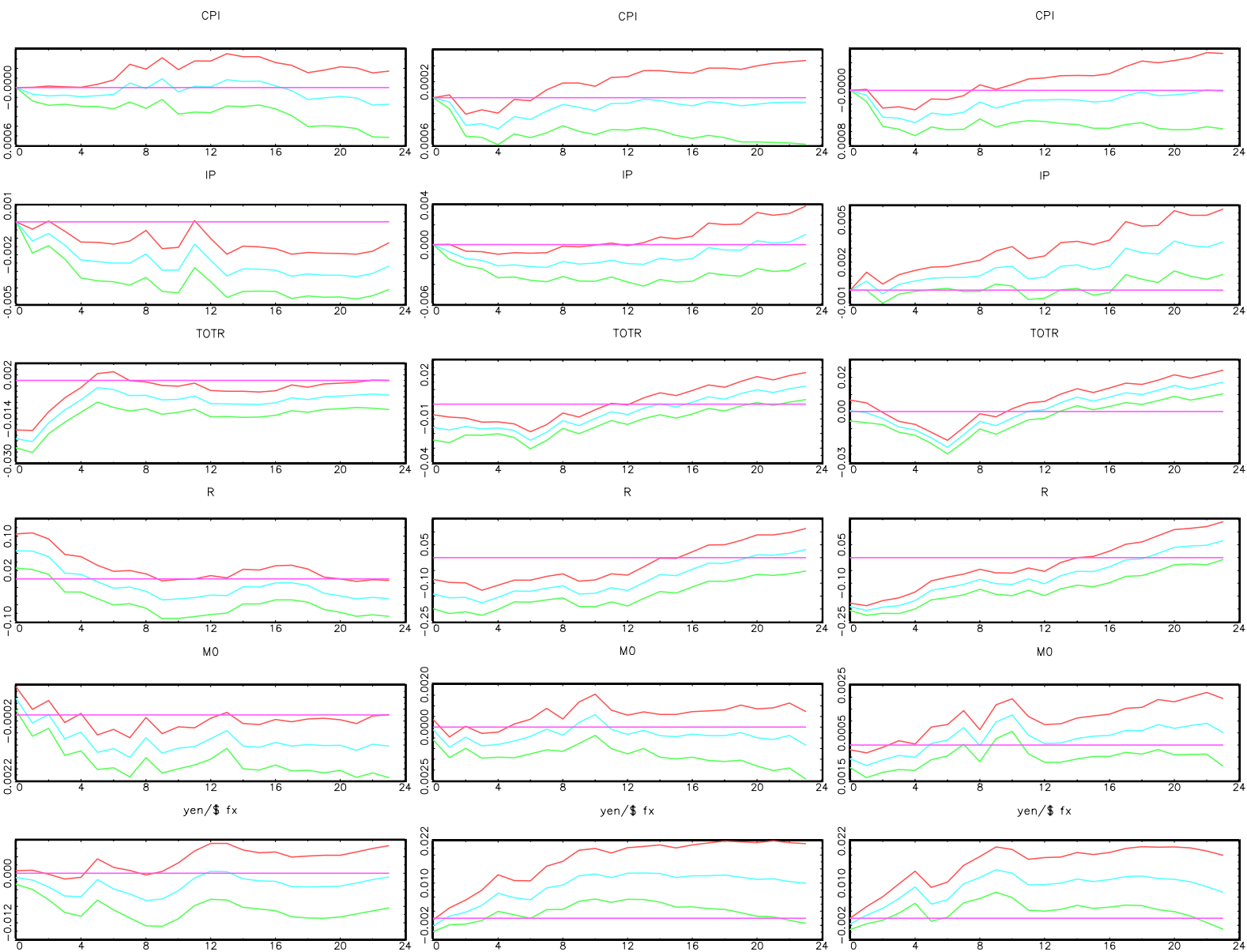
Sticky Price Model

Inflation Tax Model

Figure 5

Japan: responses of three models prices not constrained

Liquidity Effect Model

Sticky Price Model

Inflation Tax Model