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

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A cornerstone of central bank policy 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 commonly maintained in practice but are at odds with some leading models of money. This paper proposes and implements a methodology for assessing the liquidity effect hypothesis with two other hypotheses: The costly price adjustment hypothesis and the inflation tax hypothesis. We find surprisingly little support for the liquidity effect hypothesis in Japanese or U.S. data. The liquidity effect hypothesis receives its strongest support in Korean data.

Keywords: Monetary policy, Economic activity, Japan, United States

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I. 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 economists have come to agree that only surprise changes in monetary policy are likely to have these effects. But, this view about how monetary policy affects economic activity is so prevalent that many monetary economists assess the success of a model of money according to its 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 these assumptions, which we refer to as the *liquidity effect hypothesis*, are the starting point for identifying a shock to monetary policy.[1] If results from an identification scheme are inconsistent with one of these maintained assumptions, this is thought to be a shortcoming of the empirical specification. For instance if an identified expansionary monetary policy shock produces a fall in the price level it is referred to as a price puzzle and other variables such as commodity prices are included in the VAR to resolve it.

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 to be shortcoming of this class of model. Subsequent work by Lucas (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) costly price adjustment model with monopolistically competitive intermediate goods producers, interest rates and output both rise in response to a surprise increase in the growth rate of money. Christiano, Eichenbaum, and Evans (1997) provide evidence that this

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

is a robust prediction of costly price adjustment models and conclude that flexible price liquidity effect models are more consistent with the liquidity effect hypothesis.

The confidence in the liquidity effect hypothesis is so strong that it now defines the data facts used to assess empirical models of money. Christiano, Eichenbaum, and Evans (2005), for instance, assess their model's performance on the basis of the distance of model *predicted* impulse responses to monetary policy shocks from *data* impulse responses to monetary policy. The *data* impulse response functions come from an identified structural VAR that has been selected in the first place because it is consistent with the liquidity effect hypothesis.

Even though considerable efforts have been devoted to formulating theoretical models of money that are consistent with the prevailing wisdom, success has been elusive. It is surprisingly difficult to formulate either costly price adjustment or flexible price models that produce large persistent liquidity effects without appealing to *ad hoc* propagation mechanisms such as quadratic adjustment costs (see *e.g.* Basu and Kimball (2003) or Christiano (1991)) and/or assuming labor supply elasticities that are implausibly large (see *e.g.* Christiano, Eichenbaum, and Evans (1997)). Indeed, theory suggests that the liquidity effect hypothesis may not be a particularly robust phenomenon. Producing model impulse responses that are consistent with the liquidity effect hypothesis requires specific configurations of model parameters including the monetary policy feedback rule. These parameters can vary across time and countries. Moreover, producing a persistent liquidity effect depends on details of the economy that we don't have much information about.

The goal of this paper is to submit this cornerstone of modern monetary economics to more careful scrutiny and evaluate it on an equal footing with leading alternatives that are implied by theory. We use a monte-carlo procedure to 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, is the *liquidity effect hypothesis* - a surprise loosening of 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)). We will refer to these joint implications as the *costly price adjustment hypothesis*. It is important to emphasize that these hypotheses reflect economic mechanisms which may all be operating simultaneously in the same model. What we are interested in understanding is which effects are largest and thus determine the responses we see in actual data.

We evaluate each of these hypotheses by first generating monte carlo realizations that are consistent with a particular hypothesis using a 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). We impose restrictions from each of the three maintained hypotheses. We then evaluate the plausibility of each hypothesis using two metrics: A classical approach based on the likelihood function which conditions on the estimated coefficients of a reduced form VAR as is common in the structural VAR literature. We also conduct simulations that allow for parameter uncertainty in the coefficients of the reduced form VAR. This allows us to compute posterior probabilities for each maintained hypothesis under alternative sets of priors.

We find substantial empirical evidence against the liquidity effect hypothesis in data from the U.S. and Japan. For the U.S. the plausibility of the liquidity hypothesis is very sensitive to the choice of variables. If we use variables other than the ones that are known to support the liquidity effect hypothesis from the previous literature, this hypothesis is rejected. We explore the reason for these rejections and find that it lies in the output response. U.S. data is more consistent with the predictions of the costly price adjustment model which implies that the short run response of output to a higher Federal Funds rate is positive. In Japan we find that it is very also very difficult to find specifications that are consistent with the

[2] We follow Uhlig (2004) here. But see also Faust (1998) and Canova (2002) for related approaches.

liquidity effect hypothesis. The reason for the rejections in Japan is the response of prices. Removing the sign restriction on prices produces a large and persistent price puzzle. Korean data provides the most evidence in favor of the liquidity effect hypothesis. This hypothesis performs well in the period before the Asian crisis and also sample periods that include the crisis.

The remainder of the paper is organized as follows. Section II describes the theoretical motivation for the three hypotheses in more detail. Section III describes the details of our identification and evaluation procedures. Section IV contains the results and section V concludes.

II. 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 reflects the properties of a costly price adjustment model as in Rotemberg (1996). Rotemberg (1996) 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 the costly price adjustment 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 increase the number of candidates beyond the three alternatives that we consider here. Unfortunately, our understanding of how these properties of the models vary with the specification of the monetary policy feedback rule is still in its infancy (see Christiano, Eichenbaum, and Evans (2005) and Braun and Waki (2005) for recent examples of papers that relax this assumption in respectively the U.S. and Japan). Results in Aiyagari and Braun (1998) suggest that the sign responses of monetary transactions demand 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 costly price adjustment model along the lines of Rotemberg (1996). 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 most of the empirical work described above is calibrated to U.S. data. We will assume below that these hypotheses are empirically relevant for Korea and Japan, too.

III. 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.

A. The Reduced Form VAR

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

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

where x_t is a $(K \times 1)$ vector of macroeconomic variables, L is a lag operator, and $C(L) = C_1 + C_2 L + \cdots + C_J L^{J-1}$. 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

$$Px_{t+1} = PC_0 + PC(L)x_t + Pu_{t+1}, \quad E(Pu_t u_t' P') = I. \tag{2}$$

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 (2) as:

$$\tilde{x}_{t+1} = PC_0 + PC(L)P^{-1}\tilde{x}_t + \varepsilon_{t+1}. \tag{3}$$

B. 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 be consistent with the liquidity effect hypothesis. That is we want to bias things in favor of the conventional explanation about how monetary policy affects economic activity.[3] These considerations led us to consider two distinct lists of monthly variables for the U.S. and two lists of monthly variables for Korea and 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.

We use a different baseline set of variables for Korea and Japan. First, we omit non-borrowed reserves. There is no evidence that the Bank of Korea or the Bank of Japan has monitored non-borrowed reserves and in fact neither agency releases data on non-borrowed reserves. In addition, efforts to construct non-borrowed reserves from existing data in Japan have the peculiar property that they are negative for substantial sub-periods of our sample.[4] Our list of variables for Japan and Korea consists of $x_t^{JP} = (CPI_t, Y_t, TOTR_t, R_t, M0_t, FX_t)'$ where, *R* is the call 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 central banks in both Korea and Japan. In order to facilitate comparison between Korea, 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 a commodity price index. In the robustness analysis below we consider a much larger set of variables.

C. 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] 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.

[4] See Shioji (2000) for more details.

[5] We use the monthly average of the overnight rate on uncollateralized loans.

a) 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:

$$P^{-1} = \begin{pmatrix} P_{11}^{-1} & 0 & 0 \\ P_{21}^{-1} & P_{22}^{-1} & 0 \\ P_{31}^{-1} & P_{32}^{-1} & P_{33}^{-1} \end{pmatrix} \tag{4}$$

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

[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 (2003) we attempt to completely identify all of the shocks.

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

and

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

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

b) 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 (1998), 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 embody 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 IV 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 normal prior the estimated coefficients' posterior will also 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 (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 (2×2) matrix as μ_1 and μ_2, and the corresponding eigenvectors as υ_1 and υ_2. Uhlig (2001)

TABLE 1

RESPONSE OF MACROECONOMIC VARIABLES TO A CONTRACTIONARY MONETARY POLICY SHOCK UNDER THE THREE ALTERNATIVE HYPOTHESES

Hypothesis	Prices	Output	Money	Interest Rate
Inflation Tax Hypothesis	Up	Down	Up	Up
Liquidity Effect Hypothesis	Down	Down	Down	Up
Costly Price Adjustment Hypothesis	Down	Down	Down	Down

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

shows that the first column of P_{22}^{-1}, which we denote by a, has to take the following form:

$$a=\sum_{m=1}^{2}\alpha_m\cdot\sqrt{\mu_m}\cdot v_m \tag{5}$$

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

$$\sum_{m=1}^{2}\alpha_m^2=1. \tag{6}$$

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 (6). An α chosen in this way pins down the first column of P_{22}^{-1}. The second column is calculated using the restriction $P_{22}^{-1}P_{22}^{-1\prime}=\Sigma_{22}-\Sigma_{21}\Sigma_{11}^{-1}\Sigma_{21}'\equiv\Omega$. 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.

D. 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.

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 7. For money and the interest rate we will assume that the hypothesis is satisfied if the impulse response function for the respective variable has the correct sign in a majority of steps 0 through 6. 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 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 α'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 α's that satisfy all of the sign restrictions for a particular hypothesis.

The frequency of valid draws for a particular hypothesis provides information of the plausibility of a particular hypothesis. We will focus primarily on two measures of plausibility. First, we will perform simulations in which we *condition* on the estimated coefficients of a reduced form VAR and just randomize over choices of the α's. If the number of successful draws from a large number of trails is very rare we will take this as evidence against that particular maintained hypothesis.

Second, we perform simulations where we make outer-loop draws from the posterior of the estimated coefficients of the VAR as described above, then for each draw from this posterior we take multiple inner-loop draws of the α's. We then tabulate the frequency of trials for which a draw from the outer-loop yields at least one inner-loop draw that is consistent with a particular maintained hypothesis.

The frequencies tabulated in this way are used to calculate posterior odds ratios for the three hypotheses. To see how this is done, let S_i denote the $i\,th$ structure where, $\{S_i,\ i=1,2,\cdots,I\}$. A structure consists of complete specification of a data generating mechanism including the list of variables, the number of lags and the maintained hypothesis about how monetary policy affects the economy. We calculate the posterior probabilities of each structure given the data X using Bayes formula:

$$p(S_i|X) \doteq cp(S_i)p(X|S_i) \tag{7}$$

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 $p(X|S_i)$ is the probability of the data given S_i.

IV. Results

A. U.S. Data

Table 2 reports results for each of the three maintained hypotheses for two specifications the baseline specification with prices, output, non-borrowed reserves, the Federal funds rate and a commodity price index and an alternative specification where the Yen-Dollar exchange rate is used in place of the commodity price index. 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 alternatively 12, 6, and 3 lags.

Consider first the results listed under the heading of *Commodity Price*. It is standard practice in the structural VAR literature to search for an orthogonalization of the variance-covariance matrix of shocks that produces results which correspond to the liquidity effect hypothesis while conditioning on the estimated coefficients of the VAR. The results reported in the column headed *frequency of good draws* report results correspond to this same type of exercise. These simulations condition on the estimated coefficients for the reduced form VAR and then search for orthogonalizations of the variance covariance matrix of shocks that satisfy the restrictions of a particular hypothesis by randomizing over the α's. The frequency of goods draws reported in this column the fraction of 5000 random draws from the α's that satisfies the restrictions for a particular hypothesis. It is worth emphasizing that all of the results reported in this column produce the same likelihood function value. If the frequency good draws is zero as occurs in *e.g.* the case of the costly price adjustment specification with 12 lags this means that zero draws out of 5000 satisfied the restrictions of a particular hypothesis. If one takes a classical perspective to hypothesis testing, this is evidence of a rejection of that hypothesis. In this case, one would have to use a different configuration of the estimated coefficients of the VAR that is different from those given by the reduced form *unrestricted* VAR to find any successful draws and the resulting value of the likelihood function would be lower. This is because the *unrestricted* VAR estimates are MLE estimates. Since each maintained hypothesis has the same number of sign restrictions this constitutes a rejection of that specification.

Consider now the commodity price results for 12 lags reported in Table 2 under the heading *frequency of good draws*. According to the zero draw criterion the costly price adjustment model is rejected. The inflation tax hypothesis has a somewhat higher frequency of good draws than the liquidity tax hypothesis. We do *not* interpret this fact though as to say anything further about the relative plausibility of the two hypotheses. A frequency that is positive indicates that at least one orthogonalization achieves the maximum unrestricted likelihood function value. A finding that the frequency of orthogonalizations is large for a particular hypothesis says something about robustness but does not say anything about the plausibility of that hypothesis either from a classical or Bayesian perspective.

Table 2

U.S. Data

Specification:	Commodity Price		Exchange Rate	
Variables:	CPI, Industrial Production, Non-borrowed Reserves, Fed Funds Rate, Total Reserves, Commodity Prices		CPI, Industrial Production, Total Reserves, Fed Funds Rate, *M*0, $/Yen Exchange Rate	
Hypothesis	Frequency of outer-loop draws with one or more good inner-loop draws*	Frequency of good draws**	Frequency of outer-loop draws with one or more good inner-loop draws*	Frequency of good draws**
Costly Price Adjustment				
12 lags	0.434	0.000	0.566	0.022
6 lags	0.652	0.350	0.244	0.000
3 lags	0.930	0.414	0.514	0.004
Liquidity Effect				
12 lags	0.470	0.031	0.072	0.000
6 lags	0.420	0.018	0.064	0.000
3 lags	0.596	0.056	0.072	0.000
Inflation Tax				
12 lags	0.722	0.104	0.336	0.395
6 lags	0.554	0.000	0.938	0.338
3 lags	0.208	0.000	0.854	0.123

Notes: * 500 outer-loop draws from posterior of VAR and 100 inner-loop draws to decompose variance.
** Conditional on estimated VAR coefficients. 5000 random draws to decompose variance.

Next, we allow for sampling uncertainty in the estimated values of the reduced form VAR by drawing from the posterior distribution of the reduced form VAR.[7] Consider first the results for the specifications with 12 lags. This is the number of lags used in *e.g.* Christiano, Eichenbaum, and Evans (1998). If one starts from a uniform prior over the three alternative hypotheses then using equation (5) with weights of 1/3 for each hypothesis and plugging in the frequencies from column two of Table 2 for each hypothesis, the posterior probabilities for the costly price adjustment hypothesis, the liquidity effect hypothesis and the inflation tax hypothesis are respectively (0.27,0.29,0.44). These results imply that the posterior odds for the inflation tax hypothesis is about 2 relative to either of the other two hypotheses and that the costly price adjustment hypothesis and liquidity effect hypothesis have posterior odds ratios of about 1. Once we allow for parameter uncertainty there is no sense in which the liquidity effect hypothesis is more empirically relevant than the other two hypotheses.

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 one starts with prior beliefs of (0.1, 0.8, 0.1) over respectively the costly price adjustment hypothesis, the liquidity effect hypothesis, and the inflation tax hypothesis, then these results imply posterior probabilities of (0.088, 0.77, 0.147). An individual with strong priors would continue to be very confident in the liquidity effect hypothesis after viewing the evidence for the VARs with 12 lags presented in column 1.

The impulse response functions and one standard error confidence intervals for the baseline results with 12 lags are reported in column 1 can be found in Figure 1. The impulse responses are averages across good draws with 500 outer-loop replications and 100 inner-loop replications. Confidence intervals are based on standard errors for good draws. Looking first at the results for the liquidity effect hypothesis in the first column we see that the results are broadly

[7] Although discussed above, it is worth repeating that the posterior we are drawing values of VAR coefficients from does not reflect the imposition of any restrictions from a maintained hypothesis. These restrictions get imposed by throwing out bad draws.

Response of VAR variables to
shock to monetary policy:
Liquidity effect maintained hypothesis
500 outer-loop draws, 100 inner-loop draws
baseline variables

Response of VAR variables to
shock to monetary policy:
Costly price adjustment hypothesis
500 outer-loop draws, 100 inner-loop draws
baseline variables

Response of VAR variables to
shock to monetary policy:
Inflation tax maintained hypothesis
500 outer-loop draws, 100 inner-loop draws
baseline variables

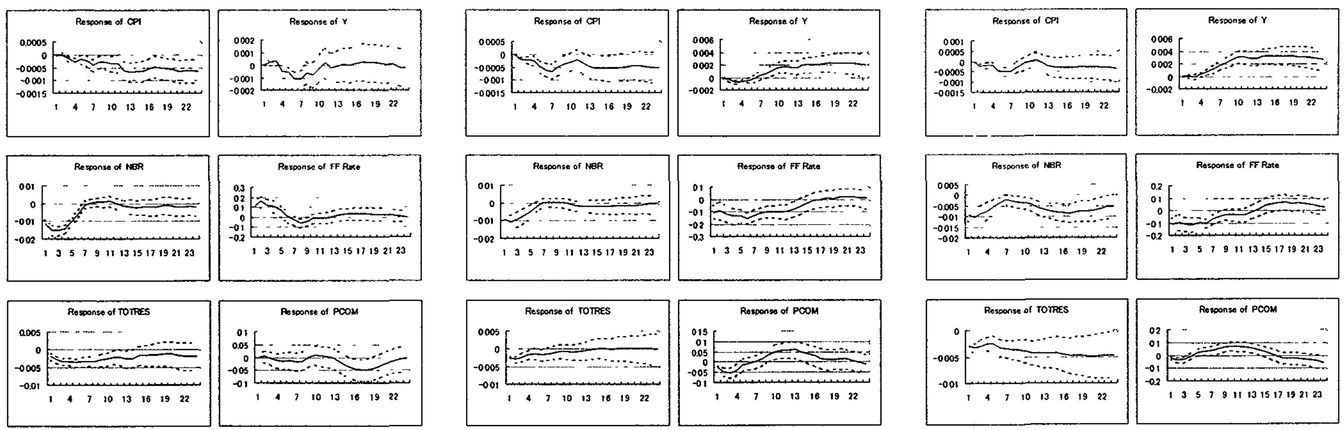

FIGURE 1

U.S. DATA IMPULSE RESPONSES PCOM SPECIFICATION, 12 LAGS 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 costly price adjustment hypothesis and the inflation tax hypothesis with 12 lags 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 costly price adjustment 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 costly price adjustment 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 costly price adjustment 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.

a) Robustness

Much of the previous VAR literature has assumed up front that the liquidity effect hypothesis is correct. The results presented so far show some evidence against this hypothesis using monthly U.S. data.

If one assigns equal prior probabilities to each hypothesis the inflation tax model receives most posterior weight. Still, a skeptic who is reasonably firm in the belief that the liquidity effect hypothesis is correct would 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. We now turn to investigate whether the conclusions are robust to variations in the number of lags and the variables that appear in the VAR.

Consider first the results for lag lengths of 3 and 6 reported in Table 2 under the *Commodity Price* heading. Our previous conclusions about the liquidity effect hypothesis are robust to the choice of lag length. We fail to reject this hypothesis using the classical criterion. Shorter lags alter the performance of the other two models though. There are now rejections of the inflation tax hypothesis. We now fail to reject the costly price adjustment hypothesis. For both the three lag and six lag specifications the costly price adjustment hypothesis now has the highest frequency of outer-loop draws with one or more good inner-loop draws. For the three lag specification if we start with a diffuse prior, the posterior odds are (0.536, 0.344, 0.120) and for a skeptical prior of (0.1, 0.8, 0.1), the posterior odds are (0.158, 0.808, 0.035). A skeptic would still assign posterior odds of about 5 to 1 in favor of the liquidity effect hypothesis over the costly price adjustment hypothesis after viewing this evidence.

Columns 3 and 4 of Table 2 under the heading *Exchange Rate* provide evidence on how the results vary with the particular choice of variables. These results use the consumer price index, industrial production, total reserves, $M0$ and the \$/Yen exchange rate in the VAR. The most striking feature of these results is that our previous conclusions about the liquidity effect hypothesis are now overturned. For this set of variables the liquidity effect hypothesis is rejected on the basis of the frequency of good draws for all choices of lag lengths. The costly price adjustment model is also rejected when the number of lags is 6. Now, if one starts with a skeptical prior of (0.1, 0.8, 0.1) the posterior distribution over the three hypotheses with 6 lags is: (0.144, 0.302, 0.553) with the inflation tax hypothesis now receiving most posterior weight.

The fact that the conclusions that one draws by looking at the two panels in Table 2 are so different raises two questions. The first question is; whether there is something special about the choice of variables used in the second panel? Our rationale for this choice of variables is that they make sense for Japan and Korea and are also available for all three countries. However, it is possible that this is a bad choice of variables. To investigate this possibility we experimented with alternative choices of variables and also changed the restrictions. Starting from the choice of variables in the right panel we tried using the commodity price index in place of the exchange rate and rejected the liquidity effect hypothesis at all lag lengths. We then tried putting the commodity price index in the first block of equations and ordered it third. The liquidity effect hypothesis was once again rejected for all lags. We then tried using $M1$ in place of $M0$ and again rejected the liquidity effect hypothesis at all lag lengths.

The second question is; what is responsible for these rejections? It is well known that the price puzzle can sometimes arise even if commodity prices are included in the VAR in U.S. data so we also tried changing the identifying restrictions and required instead that the response of prices be negative in a majority of periods 6 through 11. We continued to reject the liquidity effect hypothesis in all of the experiments we performed above. We next tried a series of runs placing non-borrowed reserves in place of total reserves and we continued to reject this hypothesis. We also repeated the above experiments by considering variants of the variables in the left panel. We reject the liquidity hypothesis if we replace the commodity price index with the foreign exchange rate. We also reject the liquidity hypothesis if we replace total reserves with a broader measure of money *e.g.* $M0$ or $M1$. Finally, we tried removing the constraint on prices and continued to reject the liquidity effect hypothesis. It is worth emphasizing that these variations are all informative in that *at least* one of the other two specifications always fails to be rejected.

Further investigations revealed that the source of the rejections is due to the response of output. We noted above that the average response of output reported in Figure 1 is positive for the liquidity effect hypothesis for the first several periods even though on a draw by draw basis the response is constrained to be negative for a majority of periods 2 through 7. This suggests that the challenge under the maintained liquidity effect hypothesis may be producing sensible output responses. To explore this possibility we took the exchange rate specification and reduced the number of periods that the sign constraint on output binds to 3 months out of the first 6 and then to 2 months out of the first 6. Under this final restriction the liquidity effect hypothesis is no longer rejected. We found similar results for the other specifications.

These results suggest that a particular choice of variables is very important for producing empirical specifications that are consistent with the liquidity effect hypothesis. If one deviates from this list of variables, the liquidity effect hypothesis is rejected in U.S. data. The source of this rejection is not a price puzzle but instead the response of output. This output response is not a challenge for the other two theories and this is why they are not rejected.

B. Results for Japan

For Japan we consider two specifications: A VAR that includes the CPI less food, industrial production, total reserves, the call rate, $M0$ and a commodity price index and a second specification that includes the Yen/$ exchange rate instead of the commodity price index. 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 using 3, 6, and 12 lags in the VAR are reported in Table 3.

Consider first the results for the specifications with commodity prices. When we condition on the estimated coefficients of the unrestricted VAR we find zero good draws for the liquidity effect maintained hypothesis for the 3, 6, and 12 lag specifications. According to this criterion the inflation tax specification with 12 lags is also rejected. When one allows for parameter uncertainty a skeptic with priors over the costly price adjustment hypothesis, liquidity effect hypothesis and inflation tax hypothesis of (0.1, 0.8, 0.1) assigns posterior probabilities of (0.33, 0.56, 0.11) to the three hypotheses. For the 3 lag specification there are no successful draws when one draws 500 times from the posterior distribution of VAR coefficients.[8] The other noteworthy feature of the results is that the performance of the inflation tax model improves as the number of the lags is

[8] We increased the outerloop draws to 5000 in order to get some good draws and found that the frequency of success was still zero to two digits under the liquidity effect maintained hypothesis.

TABLE 3
JAPANESE DATA

Specification:	Commodity Price		Exchange Rate	
Variables in VAR:	CPI, Industrial Production, Total Reserves, Fed Funds Rate, *M0*, Commodity Price		CPI, Industrial Production, Total Reserves, Fed Funds Rate, *M0*, Yen/$ Exchange Rate	
Hypothesis	Frequency of outer-loop draws with one or more good inner-loop draws*	Frequency of good draws**	Frequency of outer-loop draws with one or more good inner-loop draws*	Frequency of good draws**
Costly Price Adjustment				
12 lags	0.990	0.548	0.986	0.562
6 lags	0.976	0.488	0.908	0.319
3 lags	0.948	0.212	0.850	0.130
Liquidity Effect				
12 lags	0.212	0.000	0.294	0.000
6 lags	0.272	0.000	0.064	0.000
3 lags	0.000	0.000	0.000	0.000
Inflation Tax				
12 lags	0.340	0.000	0.430	0.026
6 lags	0.638	0.021	0.600	0.008
3 lags	0.730	0.049	0.670	0.044

Notes: * 500 outer-loop draws from posterior of VAR and 100 inner-loop draws to decompose variance.
** Conditional on estimated VAR coefficients. 5000 random draws to decompose variance.

reduced. Overall. though the costly price adjustment model is best.

The results for the specification with the exchange rate reported in Table 3 are qualitatively similar. The liquidity effect hypothesis is rejected if one applies a classical test that conditions on the data. However, if one allows for parameter uncertainty the posterior odds for a skeptic still favor the liquidity effect hypothesis when the number of lags is 12 or 6.

We report impulse responses for the 12 lag specifications with exchange rates in Figure 2. A comparison of the second and third columns in Figure 3 reveals an important difference between the costly price adjustment and inflation tax hypothesis results. The output responses under the two hypotheses are quite different. Under the costly price adjustment 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. Overall, the results shown here are consistent with the perspective that policy induced increases in interest rates lead to appreciation of the home currency.

a) Robustness

The results reported in Table 3 depend crucially on the sign restriction on prices. If this restriction is not imposed two things happen; a large and persistent price puzzle arises and the liquidity effect hypothesis is no longer rejected. One question that we explored was; could an alternative price variable or ordering resolve the price puzzle for Japan? We tried several alternative measures of prices including oil prices, a wholesale price index and also tried ordering the price variable third in the first block. We continued to get zero successful draws for the liquidity effect hypothesis for all specifications but one. If oil prices are included in a VAR with 3 lags and are ordered third then we get one successful draw out of 5000 replications when conditioning on the coefficients of the estimated VAR.

McCallum (1994) has argued that the spread on long and short

Response of VAR variables to
shock to monetary policy:
Liquidity effect maintained hypothesis
500 outer-loop draws, 100 inner-loop draws
baseline variables

Response of VAR variables to
shock to monetary policy:
Costly price adjustment hypothesis
500 outer-loop draws, 100 inner-loop draws
baseline variables

Response of VAR variables to
shock to monetary policy:
Inflation tax maintained hypothesis
500 outer-loop draws, 100 inner-loop draws
baseline variables

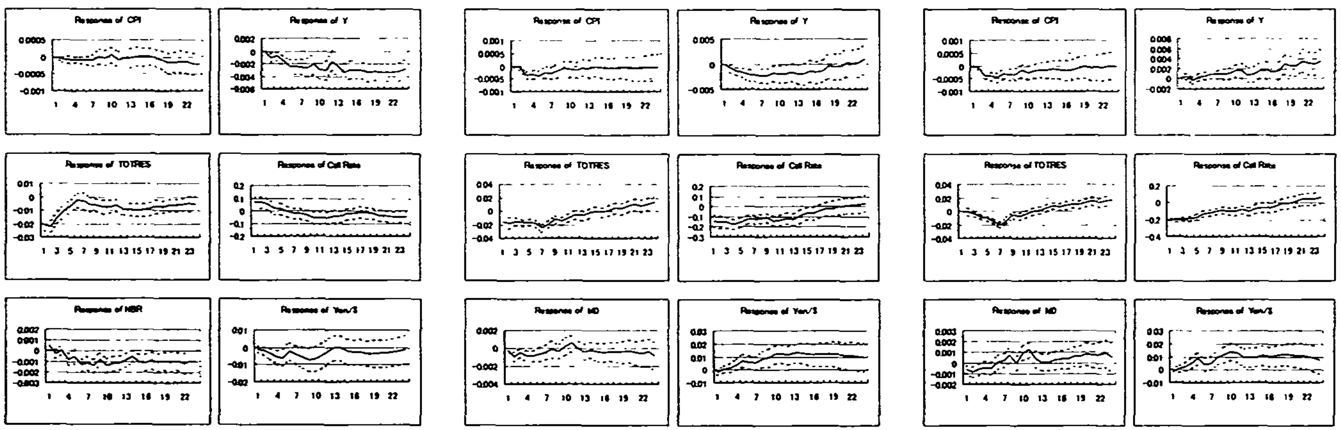

FIGURE 2

JAPANESE DATA

IMPULSE RESPONSES EXCHANGE RATE SPECIFICATION, 12 LAGS 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[9] and the Yen/Dollar exchange rate. For a specification with 12 lags the posterior distribution associated with our skeptical prior was (0.04, 0.80, 0.6). 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.

Braun and Shioji (2005) report results using a different sample period and additional yield curve variables. Simulations reported there also reject the liquidity effect hypothesis using the frequency of good draws criterion.

Taken together these results provide considerable evidence in Japanese data against the liquidity effect hypothesis and indicate further that Japanese data is most consistent with the costly price adjustment hypothesis.

C. Korean Data

Next we turn to consider results from Korea which are reported in Table 4. We consider the same two specifications that we used for Japan. The first uses monthly data on the CPI less food, industrial production, total reserves, the call rate, monetary base and a commodity price index. The second specification uses the Won-Dollar exchange rate in place of the commodity price index. The primary source of the data is the Bank of Korea homepage. In cases, where the complete time-series were not available *e.g.* Won-Dollar exchange rate, we used data from International Financial Statistics from the IMF. The sample period for Korea is chosen based on two considerations. First, the deregulation of financial markets and move towards using open market operations to implement monetary objectives is relatively recent to Korea. Deregulation of financial markets started in the early 1990s and continued throughout the 1990s. We start our sample in 1991. An earlier start date would mean including periods when interest rates were regulated and a later start date makes the sample too short to make any meaningful

[9] 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.

TABLE 4

KOREAN DATA

Specification:		Commodity Price		Exchange Rate Unrestricted		Exchange Rate Restricted	
Variables:		CPI, Industrial Production, Total Reserves, Call Rate Rate, *M*0, Commodity Price		CPI, Industrial Production, Total Reserves, Call Rate Rate, *M*0, Won/$ Exchange Rate		CPI, Industrial Production, Total Reserves, Call Rate Rate, *M*0, Won/$ Exchange Rate	
Hypothesis	Sample Period	Frequency of outer-loop draws with one or more good inner-loop draws*	Frequency of good draws**	Frequency of outer-loop draws with one or more good inner-loop draws*	Frequency of good draws**	Frequency of outer-loop draws with one or more good inner-loop draws*	Frequency of good draws**
Costly Price Adjustment							
6 lags	96:1-04:5	0.238	0.000	0.282	0.000	0.174	0.000
3 lags	96:1-04:5	0.138	0.000	0.122	0.000	0.092	0.000
3 lags	96:1-96:3	0.732	0.132	0.852	0.387	0.834	0.387
Liquidity Effect							
6 lags	96:1-04:5	0.678	0.193	0.598	0.180	0.042	0.000
3 lags	96:1-04:5	0.742	0.164	0.782	0.381	0.090	0.000
3 lags	96:1-96:3	0.994	0.605	0.998	0.383	0.318	0.000
Inflation Tax							
6 lags	96:1-04:5	0.824	0.464	0.882	0.524	0.404	0.052
3 lags	96:1-04:5	0.858	0.481	0.858	0.476	0.298	0.035
3 lags	96:1-96:3	0.004	0.000	0.006	0.000	0.002	0.000

Notes: * 500 outer-loop draws from posterior of VAR and 100 inner-loop draws to decompose variance.

** Conditional on estimated VAR coefficients. 5000 random draws to decompose variance.

inferences. The second consideration in choosing our sample period is the Asian crisis. This is a big economic event for Korea and whether one includes this period has important implications for the inferences that we draw. For these reasons, results will be reported for a sample that runs until May 2004 and thus includes the crisis period and also a shorter sub-sample that ends in March 1996.

Finally, the shortness of the sample period and a concern about the small number of degrees of freedom led us to restrict attention to VAR's with alternatively 3 and 6 lags for the longer sample period and to 3 lags for the shorter sample period.

Consider first the results using the commodity price index reported in the left panel of Table 4. For the long sample period there is considerable evidence against the costly price adjustment hypothesis. When we condition on the estimated coefficients of the VAR and perform classical inference we reject the costly price adjustment specification for the whole sample period. This is not particularly surprising given that the Asian crisis was accompanied by sharply higher interest rates, higher prices and low levels of economic activity. These same events favor the inflation tax specification. We also fail to reject the liquidity effect hypothesis for the whole sample period. Posterior odds ratios for the whole sample assign somewhat more weight to the inflation tax specification but the liquidity effect hypothesis also receives significant posterior weight. With a diffuse prior of 1/3 for each hypothesis the posterior probabilities for the costly price adjustment hypothesis, liquidity effect hypothesis and inflation tax hypothesis are respectively (0.08,0.43,.49) for the 3 lag specification.

If attention is limited to the pre-crisis sub-sample the inflation tax specification is rejected. The frequency of good draws under this hypothesis is zero. Moreover, the frequency of outer-loop draws with one or more good inner-loop draw is also very small. Under a uniform prior the posterior probabilities are (0.580, 0.418, 0.002). It is clear from this that the Asian crisis is largely responsible for the success of the inflation tax hypothesis in the longer sample period.

Using the other set of variables yields similar results. Results reported under the heading *Exchange Rate Unrestricted* show rejections of the costly price adjustment hypothesis for the longer sample period and rejections of the inflation tax hypothesis for the shorter sample period. The liquidity effect hypothesis is not rejected for either sample period.

Figure 3 reports impulse responses for the three hypotheses for the whole sample period using VARs with 3 lags and the exchange rate specification. These figures show some other evidence that favors the liquidity effect hypothesis. For the costly price adjustment specification the responses of the call rate and the CPI are both non-monotonic. These two variables switch signs as soon as the sign restrictions cease to bind. These figures in conjunction with the rejections found in Table 3 suggest there is a lot of information in the Asian crisis and that this information is strongly at odds with the costly price adjustment hypothesis. Based on the responses reported in Figure 3, it is much easier to reconcile the facts from the Asian crisis with the other two hypotheses.

Another noteworthy feature of Figure 3 is that all three hypotheses are inconsistent with the view that policy induced increases in the call rate lead to a nominal appreciation of the home currency. One possibility is that the Asian crisis is producing large movements in the exchange rate during this period that are mistakenly being attributed to monetary policy. The Asian crisis was a period where the call rate in Korea rose at the same time that the Won depreciated. If this is the case though one would expect that this puzzle would disappear in the shorter sample period. To explore this possibility consider Figure 4 which reports impulse responses for the pre-crisis period. The number of lags in the VARs is three. Interestingly, the costly price adjustment hypothesis now shows a response of the exchange rate that is consistent with uncovered interest rate parity. However, the liquidity effect hypothesis continues to produce anomalous exchange rate responses.

An advantage of our approach is that one can easily see how the answers change as one imposes more restrictions on the hypotheses. The right panel of Table 3 labeled exchange rate restricted imposes the previous restrictions plus the additional restriction that the Korean nominal exchange rate appreciate when monetary policy increases the call rate. Imposing this restriction does not affect either the costly price adjustment hypothesis or the inflation tax hypothesis but it does affect the results for the liquidity effect hypothesis. This hypothesis is now rejected on the basis of the frequency of good draws criterion for all lag/sample configurations. Imposing this additional restriction also affects the outcomes for the Bayesian test. Now the costly price adjustment specification performs best for the shorter sub-sample. Figure 5 shows how the responses for the whole

Response of VAR variables to shock to monetary policy: Liquidity effect maintained hypothesis 500 outer-loop draws, 100 inner-loop draws baseline variables

Response of VAR variables to shock to monetary policy: Costly price adjustment hypothesis 500 outer-loop draws, 100 inner-loop draws baseline variables

Response of VAR variables to shock to monetary policy: Inflation tax maintained hypothesis 500 outer-loop draws, 100 inner-loop draws baseline variables

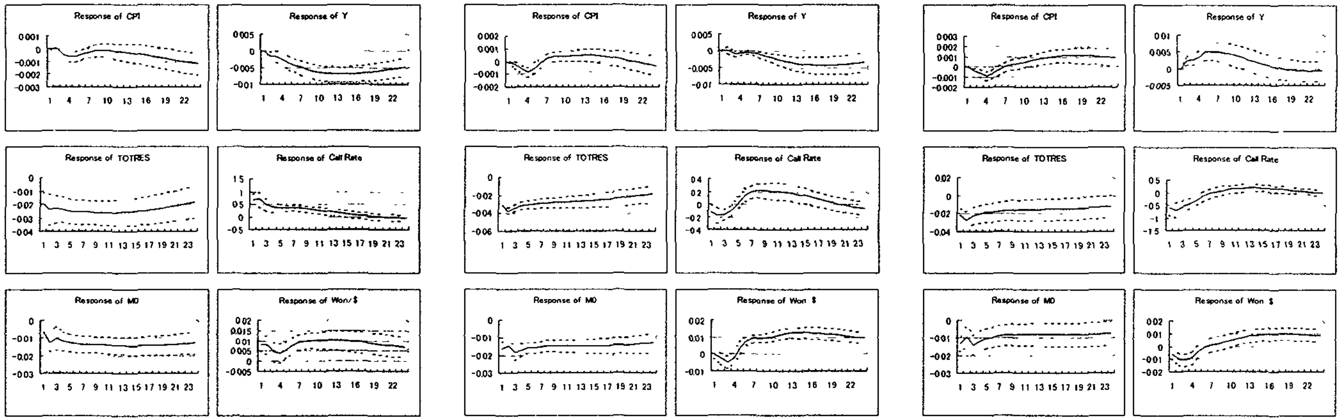

FIGURE 3

KOREAN DATA

IMPULSE RESPONSES 1991:1-2004:5 SAMPLE PERIOD

Response of VAR variables to shock to monetary policy: Liquidity effect maintained hypothesis 500 outer-loop draws, 100 inner-loop draws baseline variables	Response of VAR variables to shock to monetary policy: Costly price adjustment hypothesis 500 outer-loop draws, 100 inner-loop draws baseline variables	Response of VAR variables to shock to monetary policy: Inflation tax maintained hypothesis 500 outer-loop draws, 100 inner-loop draws baseline variables

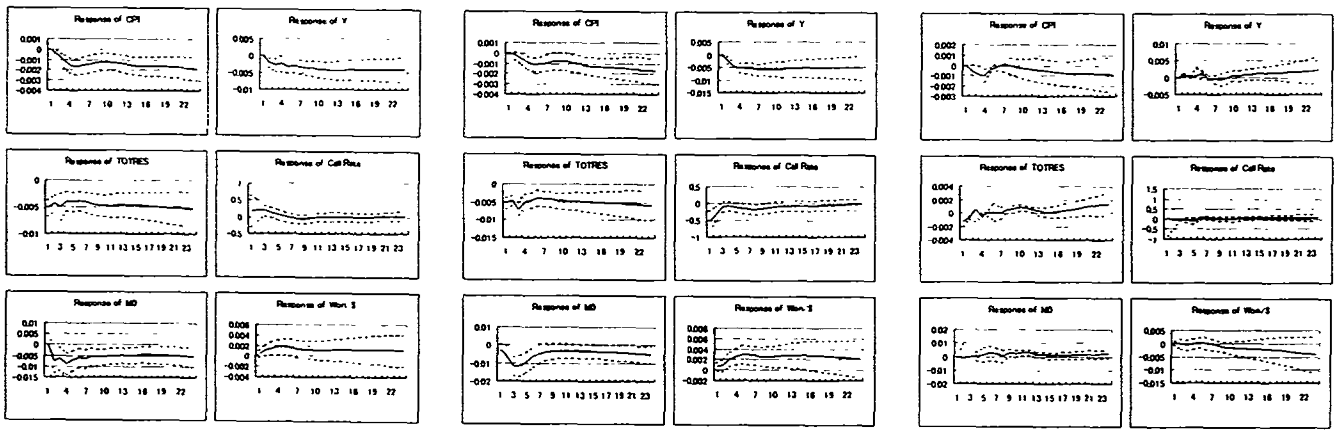

FIGURE 4

KOREAN DATA IMPULSE RESPONSES, 3 LAGS,
1991:1-1996:3 SAMPLE PERIOD

Response of VAR variables to shock to monetary policy: Liquidity effect maintained hypothesis 500 outer-loop draws, 100 inner-loop draws baseline variables

Response of VAR variables to shock to monetary policy: Costly price adjustment hypothesis 500 outer-loop draws, 100 inner-loop draws baseline variables

Response of VAR variables to shock to monetary policy: Inflation tax maintained hypothesis 500 outer-loop draws, 100 inner-loop draws baseline variables

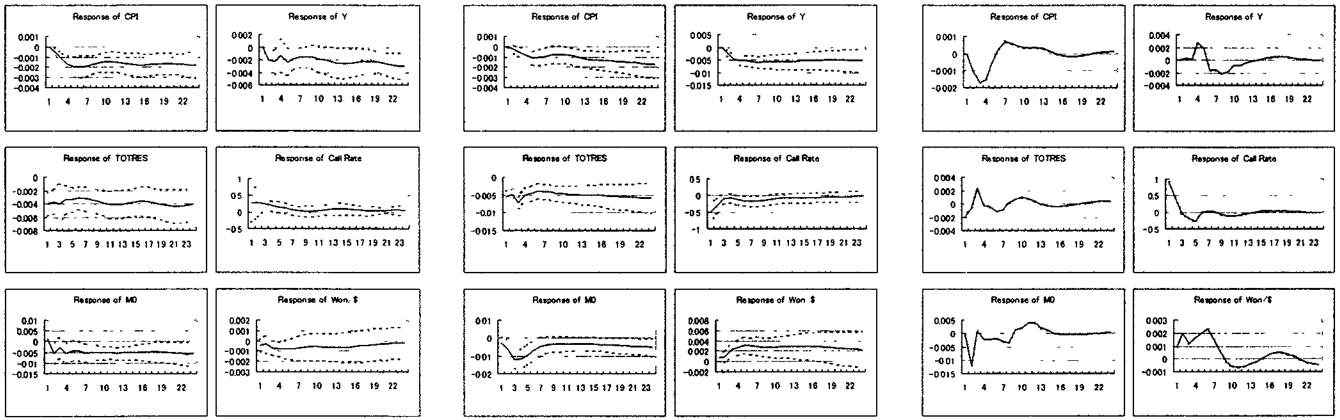

FIGURE 5

KOREAN DATA IMPULSE RESPONSES EXCHANGE RATE SIGN RESTRICTED
1991:1-1996:3 SAMPLE PERIOD

sample period are affected by imposing this further restriction.

The responses are generally similar to the previous results with the exception that the sign of the exchange rate now accords well with theory for the liquidity effect hypothesis. The costly price adjustment hypothesis still produces average impulse responses for the exchange rate that are inconsistent with uncovered interest rate parity even though the signs have been restricted for each individual draw. Moreover, the responses of the CPI and the call rate are once again non-monotonic under this hypothesis.

Although we reject the liquidity effect hypothesis on the basis of zero good draws in other respects this hypothesis lines up well with the facts. The impulse responses reported in Figure 5 look reasonable and are similar to the responses reported in Figures 3 and 4. Moreover, the posterior odds ratio under a diffuse prior for the costly price specification and the liquidity effect specification is about 1 using data from the *3rd panel* of Table 4.

As a final check on the robustness of our conclusions for Korea we also considered 3 lag VARs with the CPI, industrial production, total reserves, the call rate, *M*0 and the exchange rate in the post-Asian crisis period. The starting date is 1999:1 and the terminal date is 2004:5. The results for this sub-sample reinforce our previous conclusions that Korea data favors the liquidity effect hypothesis. If we condition on the estimated values of the VAR both the costly price adjustment hypothesis and the inflation tax hypothesis are rejected using the classical hypothesis test. However, we accept the liquidity effect hypothesis. We also accept the liquidity effect hypothesis when the additional signconstraint is imposed on the exchange rate response.

Overall, Korean data provides the most consistent evidence in favor of the liquidity effect hypothesis among the three countries we have considered.

V. 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 U.S. data we have found that if one deviates even slightly from specifications that previous research has found to support the liquidity effect hypothesis, that this hypothesis is rejected. The source of this rejection is *not* a price puzzle but instead the output response. We find it interesting that the distinction between the response of output is not a puzzle for the costly price adjustment hypothesis. Indeed, the impulse responses under this maintained hypothesis are in good accord with the predictions of theory.

We also find that the liquidity effect hypothesis is rejected for all specifications except one using monthly Japanese data. In Japan the rejection of the liquidity effect hypothesis is due to the response of prices. If the price response is not constrained, a large and persistent price puzzle arises. Posterior odds ratios generally favor the costly price adjustment hypothesis in Japan and in some cases are so large as to even convince a skeptic who assigns most prior probability to the liquidity effect hypothesis.

Korean data provides the most evidence in favor of the liquidity effect hypothesis. This hypothesis performs well in the period before the Asian crisis and also sample periods that include the crisis. The only puzzle for this hypothesis is the response of exchange rates. Unrestricted impulse responses have the property that an increase in the Korean call rate induces a nominal depreciation of the Won. This occurs both in sample periods that include the crisis and periods that pre-date it. If exchange rates are also restricted, this hypothesis is rejected using a classical test. However, posterior odds ratios and other properties of the impulse responses support this hypothesis.

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.[10] 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

[10] 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.

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.

More generally we view the empirical methodology we have described here to be an attractive way to incorporate restrictions from theory. If one is confident in a particular hypothesis our approach provides a way to impose this hypothesis on the data and assess it against other hypotheses. In addition, our approach provides a way to produce results that are robust to aspects of the empirical specification that we don't have much a priori information about such as the orthogonalization, the number of lags, and the specific choice of variables.

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