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Aaron Mehrotra

Exchange and interest rate channels during a deflationary era –

Evidence from Japan, Hong Kong and China

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Tiivistelmä

Tässä tutkimuksessa tarkastellaan rahapolitiikan korko- ja valuuttakurssikanavien merkitystä viimeaikaisten deflaatiokausien aikana Japanissa, Hongkongissa ja Kiinassa. Tarkasteltavat taloudet eroavat merkittävästi toisistaan rahapoliittisten regiimien ja talouksien avoimuuden osalta. Analyysissä käytetään rakenteellisia vektoriautoregressiivisiä (SVAR) malleja. Tulosten mukaan sokit nimellisessä efektiivisessä valuuttakurssissa vaikuttavat tilastollisesti merkittävällä tavalla hintoihin Japanissa ja Hongkongissa. Jälkimmäisessä taloudessa vaikutus on erityisen suuri. Tulokset heijastavat ulkoisten vaikutusten merkitystä hintojen kehityksessä deflaation oloissa, ja niiden voidaan myös tulkita puoltavan rahapoliittista strategiaa heikentää valuuttakurssia talouden poistamiseksi likviditeettiloukusta. Tutkimuksessa osoitetaan myös, että korkokanava on merkittävä hintoihin vaikuttava tekijä sekä Japanissa että Hongkongissa. Kiinassa, jossa korkoinstrumentilla ei ole ollut tärkeää roolia rahapolitiikan toteuttamisessa, sekä valuuttakurssi- että korkoinstrumenttien vaikutus hintoihin on pieni.

Avainsanat: deflaatio, korkojen nollaraja, rakenteellinen vektoriautoregressio (SVAR)
JEL-luokittelu: E31, F41

Exchange and Interest Rate Channels during a Deflationary Era - Evidence from Japan, Hong Kong and China*

Aaron Mehrotra

November 15, 2005

Abstract

We examine the role of the exchange and interest rate channels during recent deflation episodes in Japan, Hong Kong and China. We estimate open-economy structural vector autoregressive (SVAR) models for the three economies with different monetary regimes and varying degrees of openness. In both Japan and Hong Kong, shocks to the nominal effective exchange rate have a statistically significant impact on prices, with a notably stronger effect in Hong Kong. Our results provide evidence about the role of external influences in the deflation episodes of these economies, and could also be seen to weakly support suggestions to depreciate the currency in order to escape from a liquidity trap. The importance of the interest rate channel is also found to be high in Japan and Hong Kong. In China, where interest rates have not been an important monetary policy tool, neither exchange nor interest rate shocks significantly influence price developments.

Keywords: Deflation, Zero lower bound, SVAR

JEL Classification: E31, F41

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1 Introduction

The increased occurrence of deflation episodes in both emerging and industrial economies has engaged the concern of policymakers, calling for an assessment of the possible causes and cures for falling prices. Japan's recent stagnation-deflation episode has undoubtedly received the most headlines and produced the most calls for policy proposals from economists, but, in fact, deflation rates in Hong Kong SAR, which uses a currency board arrangement, were actually higher. Meanwhile, mainland China's bout with deflation occurred under circumstances that combined a dollar peg, capital controls, and rapid economic growth. That these three economies, with such diverse monetary regimes and macroeconomic environments, experienced fairly contemporaneous deflation episodes suggests a fertile opportunity for evaluating the potency of interest rate and exchange rate instruments in deflationary economies, with the ultimate aim of reimposing price stability or inducing modest inflation.

We thus examine the role of the exchange rate and interest rate channels during these recent Asian episodes of disinflation and deflation by estimating structural vector autoregression (SVAR) models for Japan, Hong Kong, and China. Our open-economy framework allows for the identification of structural interest rate shocks and monetary shocks originating in the exchange rate. As the Hong Kong dollar and the Chinese renminbi were pegged to the US dollar during the time of the study, we use nominal effective exchange rates in our analysis. Numerous previous studies discuss the causes of inflation and deflation in these economies; we hope to add to this knowledge through incorporating some novel elements in our approach. Using a similar estimation framework for all three economies for the same time period, we attempt to capture similarities and differences between price-adjustment processes. Further, there are to our knowledge no previous estimations of an open-economy SVAR model for a period of falling prices, allowing for the assessment of the relative importance of exchange rate and interest rate shocks.

The exchange rate tool is often suggested as a means for escaping from a liquidity trap. The best-known approach is perhaps the "Foolproof Way" proposed by Svensson (2001, 2003). In contrast, the attraction of the interest rate instrument has been predominantly seen as lost once the zero bound is hit. In particular, we note that the influential Taylor-type interest rate feedback rules (see Taylor, 1993), which characterize the interest rate setting behavior of a central bank as a linear function of inflation and the output gap, generally do not take into account the zero-floor bound for interest rates.

For both Japan and Hong Kong, we find that shocks to the nominal effective exchange rate have a significant impact on prices (with a notably stronger effect on the Hong Kong economy). Our results also give insight into the role of external influences in the deflation episodes of both economies. The fact that Hong Kong endured high nominal and real interest rates during the deflationary era also points to a possible drawback of Hong Kong's currency board arrangement, which restricts domestic interest rate policy. Indeed, the importance of the interest rate channel in price development is found non-negligible for both Hong Kong and Japan. The results for Japan support a strategy of allowing the yen to depreciate to escape the liquidity trap, even if the required exchange rate movement is substantial. For China, where interest rates have not been an important monetary policy tool, neither exchange rate nor interest rate shocks appear to significantly influence price development.

This paper is structured as follows. The following section provides a descriptive analysis of macroeconomic developments in the three economies. We next discuss theoretical issues pertinent to the research question. The econometric methodology of the study is then tackled, followed by a presentation of the empirical results and a discussion. The final section concludes.

2 Developments in the Three Economies

The following is a descriptive analysis of developments in consumer prices, interest rates, and exchange rates in Japan, Hong Kong, and China. We also mention some of the relevant literature on deflation in these economies.

The Japanese economy saw consumer price inflation fall below zero in late 1995. Positive inflation re-emerged along with meager growth in 1997, then deflation took hold again. Overall, however, Japan's consumer price index remained relatively stable throughout the period of our study (Japan, 1991Q1–2004Q2; Hong Kong, 1991Q1–2004Q3; China, 1996M1–2004M8). This is evident from Figure 1 below, where we depict the annual consumer price inflation rates for all three economies (the shorter time series in many figures for China stem from limited data availability). The deflationary episode in Japan has been broad-based as prices for major items in the consumer price index (food, clothing, transportation, and durable goods) have all registered declines (IMF, 2003). The nominal exchange rate depreciated sharply after peaking in 1995.[1] In mid-1998, the exchange rate rebounded and a new appreciation peak was hit in 2000 (see Figure 2). Ultimately, the Bank of Japan encountered the well-documented limitations of conventional monetary policies reliant on the interest rate channel.[2] A zero interest rate policy was initiated by the Bank of Japan in early 1999, briefly abandoned in August 2000, and re-introduced in March 2001. The policy rate and other short-term rates all hovered close to the zero bound (the behavior of the 3-month CD rate is depicted in Figure 3).

[1] Having appreciated 20% against the dollar between January and April 1995, the yen was in full retreat by September of that year.

[2] The movements in the yen exchange rate may have partly offset the stimulus from lower interest rates. According to Meredith (1998), an IMF analysis suggests yen appreciation limited the easing in monetary conditions after asset prices collapsed.

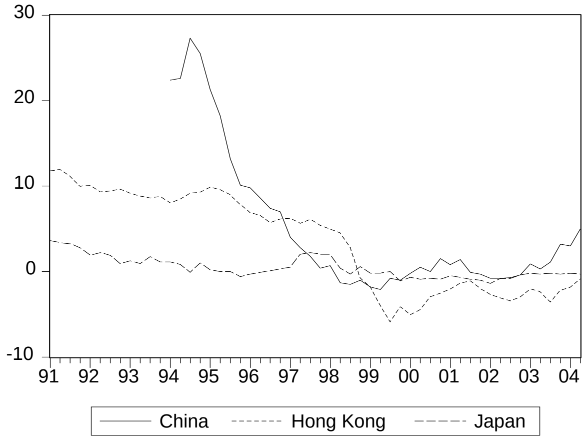

Figure 1. Consumer price inflation rates (year-on-year), China, Hong Kong, and Japan, 1991–2004.

Hong Kong, which still had CPI inflation of 10% in early 1995, saw its inflation rate plummet and deflation emerge in late 1998. By mid-1999, deflation was running at 6% a year. Thereafter, the deflation rate moderated and positive annual inflation rates reappeared in the second half of 2004. The main factors feeding deflation were prices for housing, food, durable goods, and clothing. Schellekens (2003) shows that these items together contributed to over 85% of overall deflation. His empirical study finds that deflation in Hong Kong SAR is best explained by cyclical shocks, amplified by balance-sheet and wealth effects. Further, the role of price equalization with mainland China is found to be small. Notwithstanding a fall in 1998 at the time of the Asian financial crisis, the nominal effective exchange rate of the Hong Kong dollar steadily appreciated until 2002. Thereafter, the exchange rate depreciated along with the fall in the US dollar (see Figure 2). Here, the currency board arrangement, based on the US dollar, limited the ability of the Hong Kong Monetary Authority to implement policy through the interest rate channel. Hong Kong interest rates closely track US rates, except for the turbulent period from October 1997 to end-1998 (see Figure 3). The Hong Kong Monetary Authority was confronted by massive speculation on the Hong Kong dollar in October 1997, inducing the overnight rate to reach a maximum of 280% (Gerlach, 2005).

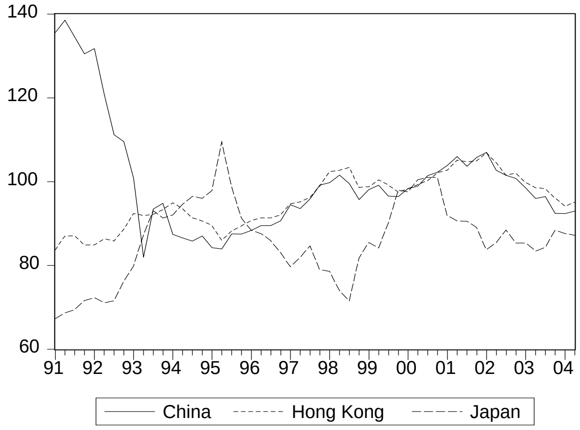

Figure 2. Nominal effective exchange rates (index, foreign-currency price of local currency). China, Hong Kong, and Japan, 1991–2004.

Chinese consumer prices started to drop in 1998 as the economy absorbed the effects of the Asian financial crisis. Deflation lasted until 2000 and reappeared in late 2001. Positive annual inflation rates returned in early 2003. Among studies of the recent inflation dynamics in the mainland China economy, Ha et al. (2003) observes that the value of the renminbi, world prices, and productivity levels were the important determinants of long-term price movements in China. Due to the US dollar peg of the Chinese renminbi during our estimation sample, the development of the nominal effective exchange rate closely resembles that of the Hong Kong dollar, whereby a long period of appreciation gives way to depreciation. The latter trend is often cited in complaints (especially US) about an undervalued renminbi. The IMF (2003) argues that transitory (lower commodity prices and WTO related tariff cuts) and long-term supply shocks (productivity gains, new technology and increased competition) were behind China's recent deflation. In a similar vein, Cargill and Parker (2004) claim that Chinese deflation was supply-led. Finally, while China's monetary policies have gradually moved toward deregulation that allows Chinese banks wider departures from official reference rates on credits and deposits, Korhonen (2004) points out Chinese interest rates have been nominally high, and rather stable, in the deflationary climate of recent years. This stable behavior is reflected in the average repurchase rate for China since mid-1999 (see Figure 3).

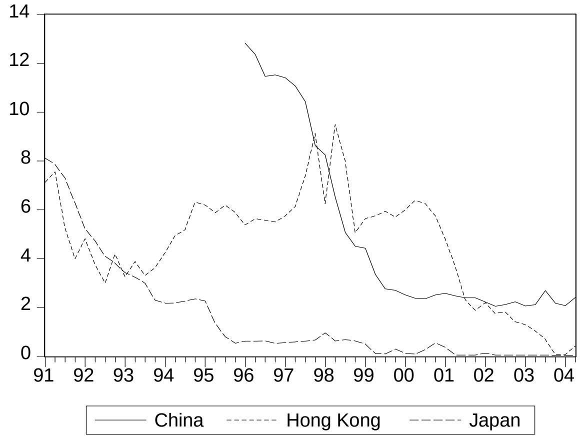

Figure 3. Short-term interest rates (3-month rate for Japan and Hong Kong, average repurchase rate for China), 1991–2004.

3 Theoretical Considerations

We next discuss theoretical considerations relevant to the research question. Japan's deflation experience and the prolonged economic stagnation accompanying it has inspired a number of studies suggesting strategies monetary authorities can use to avoid the zero interest rate floor or escape a deflationary spiral. While the latter studies likely hold more relevance for the work at hand, emphasizing the importance of the exchange rate channel, the former are also given some prominence. Even if the short-term interest rates were indeed declining throughout the data sample, a possibly even more aggressive monetary easing may have been called for before the zero bound was actually hit.

Svensson (2001, 2003) proposes a "Foolproof Way" for escaping a liquidity trap. It has three elements. First, the central bank must announce an upward-sloping target path for the domestic price level. Second, it must declare a currency devaluation and the institution of a crawling currency peg. Third, an exit strategy is stated, i.e. when a price level target path is reached, the currency peg is abandoned in favor of flexible price level targeting, at which point either the former target path is reinstituted or flexible inflation targeting based on an earlier inflation target.

It is important to note that the effect of the depreciation on the domestic price level goes beyond the increase in import prices and the stimulation of exports in the home economy. The currency depreciation and the crawling peg establish the central bank's credibility and induce expectations of higher price levels in the private sector. This impacts real ex ante interest rate and stimulates the economy. The impacts on credibility and price expectations should be advantageous as the theoretical literature on open-economy macroeconomics generally predicts a slow, rather small, exchange rate pass-through. This outcome obtains, for example, under a pricing-to-market assumption, whereby import price-setters set prices equal to those of domestic goods. Exchange rates can then simply be "disconnected" from other macro variables (see e.g. Engel, 2002).

McCallum (2000, 2003) proposes a policy rule for central banks that have already hit the zero bound. Here, the central bank must subordinate the nominal exchange rate (as opposed to the interest rate) to macroeconomic conditions. Under this rule, the rate of depreciation must increase once inflation and/or output are below target values. In place of a rules, Orphanides and Wieland (2000) focus on the effects of monetary base on aggregate demand and inflation at the zero bound. The impact emphasized by the authors is derived from the portfolio-balance effect, whereby the exchange rate reacts to changes in the relative domestic and foreign money supplies. Coenen and Wieland (2003) consider the effectiveness of the three aforementioned proposals and find the "beggar-thy-neighbor" implications of the proposals to be non-negligible.

The prevailing exchange rate regime has implications for inflation transmission across borders, and is of some interest for our three economies which have different exchange rate arrangements. Flexible exchange rate regimes are generally considered to offer the most favorable insulation properties with respect to foreign inflationary pressures by allowing for independent monetary policy (Friedman, 1953). Yet even when an economy with flexible exchange rates hits the zero interest bound, conventional interest rate policy is constrained. This may increase the attractiveness of interventions in foreign exchange markets.[3] Earlier media discussion of Chinese deflation centers on the possible transmission of falling prices in this economy to other economies in the region and the United States (see IMF, 2003). Less attention has been paid to the possibility that China itself may be affected by external deflationary pressures, especially as an appreciation in its nominal effective exchange rate was followed by domestic deflation, and the re-emergence of positive inflation rates has coincided with a period of a depreciating renminbi. Our study provides little evidence to support this, however, as the importance of exchange rate shocks in the movement of Chinese consumer prices appears to be quite small.

The attractiveness of the exchange rate instrument is strengthened when Taylor-type interest rate feedback rules are considered in the context of the zero bound. Benhabib, Schmitt-Grohé and Uribe (2002) show how application of Taylor rules can drive an economy into a liquidity trap by establishing self-fulfilling decelerating inflation paths and aggregate fluctuations. The authors demonstrate how the destabilizing properties of Taylor-type rules emerge by considering the zero bound on nominal interest rates and global equilibrium dynamics. Measures have also been considered to make the interest rate channel potent even at the zero bound. One suggested approach is to render the zero bound irrelevant by making negative interest rates feasible. This may be achieved, as suggested recently by Buiter and Panigirtzoglou (2003) and Goodfriend (2000), through a currency tax. By imposing a carrying cost on money, the central bank can achieve any negative interest rate it desired. Alternatively, Eggertsson and Woodford (2003) and Jung, Teranishi and Watanabe (2001) consider

[3] Fujii (2004) examined the policy action of depreciating the yen with the aim of fighting deflation. The author estimates the pass-through from the exchange rate to import prices, finding that the inflationary effect of depreciation actually declined substantially during the past two decades.

the optimal policy for a central bank already enmeshed in a liquidity trap. They suggest creating private sector expectations of higher future inflation, reducing real interest rates, and stimulating aggregate demand. They show that policy defined in terms of the nominal interest rates can stimulate aggregate spending even at the lower bound as long as the public expects that the central bank will keep interest rates low for a long time and will not offset future inflation.

4 Methodology

In this section, we present the methodology used in the study, introducing the data sources and the precise variables used in the estimation. The time-series properties of the individual series are analyzed, followed by a presentation of the SVAR estimation framework for the three economies. References to related work using the SVAR methodology are discussed.

4.1 About the Data

For all the three economies, we use a four-variable system comprised of indicators for real output, consumer prices, interest rates, and the nominal effective exchange rate. Frequency of the variables varies depending on data availability. For Japan, we use quarterly data for the period 1991Q1–2004Q2, while the estimation sample for Hong Kong is longer by one observation, reaching to 2004Q3. Monthly data for the period 1996M1–2004M8 is used for the Chinese economy.[4] The data sources are the OECD Main Economic Indicators, International Financial Statistics, and Thomson Datastream databases. All output and price series in levels are seasonally adjusted. A logarithmic transformation is applied to all other series, except interest rates and the Chinese year-on-year inflation rate. The data sources and transformations are detailed in Appendix A. Estimations are conducted using the software JMulTi (2004), version 3.11. All series used in the estimations are depicted in Appendix A.

We now consider the system variables. For real output, we use real GDP for Japan and Hong Kong. For China, the choice of the output variable is more problematic as quarterly real GDP series are only available from 2000 onward. Industrial production (year-on-year growth rates) series are available from 1994 onward at the IFS database, but this series finishes in October 2002. We thus use data on quarterly nominal GDP, which are available for the entire data sample, and linearly interpolate the series into monthly observations. Finally, we construct a consumer price index series (only annual inflation rates are available) to deflate the GDP series into a real output variable.[5] For

[4] Our choice of estimation sample allows for the inclusion of the disinflationary and deflationary periods for both Japan and Hong Kong, while the financial bubble in Japan of the late 1980s is excluded from the analysis. We feel it is imperative to employ monthly data for China to have an adequate number of observations in the analysis. For Japan, misspecification tests produce unsatisfactory results for long lag lengths when monthly data are used; a high VAR order is arguably necessary to adequately capture the dynamics between the variables with high frequency data.

[5] Data on cumulative year-on-year GDP growth rates are also available for the entire data sample for China. After attempting this approach in an earlier version of this paper, it was rejected. The results were hard to compare among

prices, the consumer price index is used for Japan and Hong Kong. Annual CPI inflation rates are used for China due to the identical order of integration with the other series for this economy (as discussed in the context of the time series properties of the data).

The use of the nominal effective exchange rate (NEER) for all economies is justified by the US-dollar pegs of the Chinese renminbi and the Hong Kong dollar during the time frame of our study. As a consequence of the peg, the NEER is determined completely outside Hong Kong and China, and depends on the exchange rates between these economies' trading partners and the US dollar. Moreover, the nominal effective exchange rate of the yen is well known to closely follow the US-yen rate. This is advantageous for our analysis as foreign exchange interventions to stabilize the value of the yen have largely focused on the dollar-yen rate. For interest rates, we use the 3-month certificate of deposit (CD) for Japan, the 3-month interbank rate for Hong Kong and the average repurchase rate for China.[6] In the case of China, the reference one-year lending rate is administratively set. Its level was adjusted just eight times during the estimation period.[7] Therefore, we apply the average repurchase rate, which rather closely tracks the one-year lending rate. Figure 4 displays the reference lending rate and the repurchase rate used in our study. The coefficient on contemporaneous correlation between the series is found to be high, 0.98.

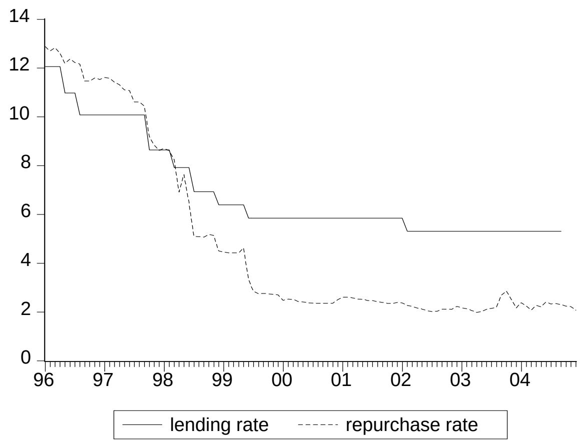

Figure 4. One-year reference rate for lending and average repurchase rate. China, 1996–2004.

the three economies.

We acknowledge the limitation of the linear interpolation for the Chinese GDP series in that this procedure does not add to the available output information for this economy. There is, however, a lack of suitable series (e.g. industrial production) for information about monthly movements. It is also not clear that other methods of output data construction would have produced preferable GDP series, given the arguments about quality problems with official Chinese GDP data (see Rawski, 2001).

[6] None of these rates are the policy rates of the central bank. However, we argue that our chosen variables may better illustrate the cost of lending for households and businesses in Japan and Hong Kong. Moreover, in Japan with the emergence of the zero interest rate policy in 1999, there was no further change in the uncollaterized overnight call rate (except for the short break in the zero rate policy) even if the CD rate continued to fluctuate, albeit at very low levels, until early 2001.

[7] The PBoC reacted to the slowdown of 1996-1998 by lowering administratively set nominal interest rates five times in the period (Roberts and Tyers, 2003).

The impact of exchange or interest rates on prices may be analyzed in a single equation framework, for example, by estimating a simple AS curve (e.g. a New Keynesian Phillips curve, where the nominal effective exchange rate enters the real marginal cost variable). A vector autoregressive (VAR) approach, on the other hand, allows for simultaneous examination of interest rate and exchange rate shocks. As it is sometimes claimed that it is difficult to give economic interpretations to reduced form VAR equations, this problem is avoided by using structural vector autoregressions in which economic theory or econometric considerations are used to impose the structure of the model. In this class of models, identification focuses on the errors of the system, which, in turn, are interpreted as linear combinations of exogenous shocks.

To determine the order of integration of the underlying series, Augmented Dickey-Fuller (ADF) tests are performed for all the series (complete results listed in Appendix B). The tests give rather strong evidence that all the series in levels, together with the inflation rate for China, are integrated of order one.[8] Due to this order of integration, it is possible that the series share common stochastic trends, i.e. cointegrating relationships. Testing for such relations in our short sample for Hong Kong and Japan is not likely to yield reliable estimates, however.[9] Moreover, if the system is estimated in levels specifying a lag length of 2 or higher, the usual tests and t-values have their asymptotic properties (Dolado and Lütkepohl, 1996) and the estimation framework would not be misspecified.[10] Our choice then is to estimate the model in levels for Japan and Hong Kong. For China, we opt for estimating a model in first differences. This procedure yields more satisfactory results in terms of the misspecification tests, as well as reflects our low enthusiasm for analyzing cointegration (long-run) relationships at a time when the Chinese economy is experiencing deep structural change. Moreover, when the Saikkonen-Lütkepohl cointegration test (see Saikkonen and Lütkepohl, 2000) is used in testing for common stochastic trends for China, the null hypothesis of no cointegration cannot be

[8] In the case of the interest rate series for Japan, the ADF test provides evidence that the series in levels were stationary. An alternative test for unit roots, the KPSS test proposed by Kwiatkowski et al. (1992), suggests a rejection of the null hypothesis of trend stationarity for Japan for the series in levels at a 1% level when 2, 3 or 4 lags are used in the test. The Japanese price level is also found to be integrated of order one in the unit root test with a structural break for 1997Q2 (see Lanne et al., 2002), corresponding to the consumption tax hike at the time. In the case of Hong Kong, a unit root test with a structural break provides evidence that the price series in first differences is stationary. This result is attained when a shift dummy is included in the system, obtaining a value of zero before 1998Q3 and one thereafter. This date corresponds to the Asian crisis and the rapid drop in the rate of inflation at that time. We then continue with the assumption that all the series in levels, together with the annual inflation rate for China, are integrated of order one, despite the fact that including a shift dummy in the VAR system in levels for Hong Kong may not guarantee that the price level series is actually I(1) for this economy.

[9] If the maximum lag length in the Saikkonen-Lütkepohl test (see Saikkonen and Lütkepohl, 2000) is set to 4 (taking into account the extremely short sample size), a cointegration rank of zero cannot be rejected for Japan at any conventional significance level for any lag length afforded by the various information criteria. This result holds when a constant and trend are included as the deterministic terms, and all the system variables are included in the test. For Hong Kong, an identical result holds for the lag lengths yielded by the Akaike and Hannan-Quinn information criteria, whereas using the Schwarz information criteria suggests the existence of a cointegration rank of one for this economy.

[10] A similar argument holds for imposing long-run restrictions in order to identify the structural vector autoregressive model. In particular, it is difficult to distinguish demand from supply shocks when there are limited signs of the economies experiencing cycles of any kind, e.g. the Japanese economy slowed for most of the estimation sample, while high growth rates persisted in the Chinese economy.

rejected at conventional significance levels for any lag length suggested by the various information criteria, including a constant and trend as the deterministic terms. These results are consistent with tests on trivariate subsystems as shown in Appendix C.

Taking the previous considerations into account, we opt for the estimation of a structural VAR model with contemporaneous restrictions and ignoring possible cointegration relations. This approach is common in the literature. Cushman and Zha (1997) apply a similar methodology for Canada, a small open economy under flexible exchange rates. Kim (1999) examines the G-7 group of countries. Kim and Roubini (2000) focus on the G-7, excluding the US. In contrast, Jang and Ogaki (2003) use long-run restrictions in a structural vector error correction (SVEC) model to study the impacts of Japanese monetary policy shocks on exchange rates. None of the studies mentioned, however, examine a deflationary period or an economy close to or at the zero bound. The estimation period for Japan in Kim and Roubini (2000) ends in 1992, the period in Jang and Ogaki (2003) in 1993, and Kim (1999) extends through 1996. To our knowledge, a structural VAR model with contemporaneous restrictions has yet to be used for Hong Kong or China to examine the effects of monetary policy during a deflationary period.

4.2 The SVAR Approach

A reduced form VAR model can be written as in Lütkepohl (2004):

$$x_t = A_1 x_{t-1} + ... + A_p x_{t-p} + CD_t + u_t \tag{1}$$

where p denotes the order of the VAR model. Here, K is the number of variables, $x_t = (x_{1t}, ..., x_{Kt})'$ is a $(K \times 1)$ random vector, A_i are fixed $(K \times K)$ coefficient matrices and D_t is a vector of deterministic terms. C is the coefficient matrix associated with the possible deterministic terms, such as a constant and a trend. The $u_t = (u_{1t}, ..., u_{Kt})'$ is a K-dimensional white noise process with $E(u_t) = 0$. Ignoring the deterministic terms, a structural representation of (1) can be expressed as:

$$\mathrm{A} x_t = A_1^* x_{t-1} + ... + A_p^* x_{t-p} + \mathrm{B}\varepsilon_t \tag{2}$$

where $\varepsilon_t \sim (0, I_K)$. Here, the matrix A allows for the modeling of instantaneous relations. The A_i^*'s $(i = 1, ..., p)$ are $(K \times K)$ coefficient matrices, while B is a structural form parameter matrix. The structural shocks, ε_t, are related to the model residuals by linear relations. Note deterministic terms are omitted; they are unaffected by impulses hitting the system and do not affect such impulses. Structural shocks are assumed to be mutually uncorrelated and therefore orthogonal. Because it is impossible to directly observe the structural shocks, certain assumptions are necessary for identification. The connection between the reduced and structural form is obtained simply by multiplying (2) with A^{-1} , with $A_j = \mathrm{A}^{-1} A_j^* (j = 1, ..., p)$. The relation between the reduced-form disturbances and the structural-form innovations are then expressed as in Breitung et al. (2004):

$$u_t = \mathrm{A}^{-1}\mathrm{B}\varepsilon_t \tag{3}$$

We estimate the AB-model of Amisano and Giannini (1997). In this case, the model for innovations can be written as Au_t =Bε_t. Linear restrictions on A are written in explicit form as vec(A)= $R_{\mathrm{A}}\gamma_{\mathrm{A}}$ $+r_{\mathrm{A}}$, where γ_A contains all unrestricted elements of A, R_{A} is a suitable matrix with 0-1 elements, and r_{A} is a vector consisting of zeros and ones. Similarly, linear restrictions on B are expressed as vec(B)= $R_{\mathrm{B}}\gamma_{\mathrm{B}} + r_{\mathrm{B}}$. Together, these two sets of restrictions are used to identify the system, i.e. matrices A and B. The number of non-redundant elements of Σ_u, $K(K+1)/2$, is the maximum number of identifiable parameters in matrices A and B. The overall number of elements in the structural form matrices A and B is $2K^2$. We therefore impose $2K^2-\frac{K(K+1)}{2}$ further restrictions to identify the full model.

Our system includes four endogenous variables: y_t (real output), p_t (prices), i_t (interest rate) and $neer_t$ (exchange rate). The errors of the reduced form VAR are written as $u_t = (u_t^y, u_t^p, u_t^i, u_t^{neer})'$. The structural disturbances $\varepsilon_t^y, \varepsilon_t^p, \varepsilon_t^i, \varepsilon_t^{neer}$ are output, CPI, monetary policy (interest rate), and exchange rate shocks, respectively. The AB-model in the form Au_t =Bε_t is written in the case of Japan as:

$$\begin{bmatrix} 1 & 0 & 0 & 0 \\ a_{21} & 1 & 0 & 0 \\ 0 & a_{32} & 1 & 0 \\ a_{41} & a_{42} & a_{43} & 1 \end{bmatrix} \begin{bmatrix} u_t^y \\ u_t^p \\ u_t^i \\ u_t^{neer} \end{bmatrix} = \begin{bmatrix} b_{11} & 0 & 0 & 0 \\ 0 & b_{22} & 0 & 0 \\ 0 & 0 & b_{33} & 0 \\ 0 & 0 & 0 & b_{44} \end{bmatrix} \begin{bmatrix} \varepsilon_t^y \\ \varepsilon_t^p \\ \varepsilon_t^i \\ \varepsilon_t^{neer} \end{bmatrix} \tag{4}$$

As $2K^2-\frac{K(K+1)}{2}$ restrictions are needed for exact identification (22 in the case of four endogenous variables), this model is overidentified. This can be easily verified by adding up the number of zeros and ones in the above matrices. However, as shown later in the context of model estimation, a formal test for over-identification is not rejected by the data. In our model, as in Kim and Roubini (2000), real activity responds to price and financial signals (both exchange and interest rates) with a lag. Moreover, prices react contemporaneously to real activity only, while their reaction to monetary policy and exchange rate shocks takes place with a lag. The adjustment lag of real output and prices to monetary shocks is a canonical assumption in monetary policy analysis. The non-contemporaneous impact of the exchange rate on prices is further justifiable on the basis of empirical evidence of a slow pass-through from the exchange rate to consumer prices.[11] The third row in the system above can be read as a monetary policy reaction function, whereby the interest rate is set solely on the basis of the price level and does not contemporaneously react to the exchange rate. As argued by Smets (1997), such behavior is characteristic of an interest rate-targeting regime that largely neglects movements in the exchange rate in its policy. This well characterizes the behavior of Japan's monetary authority in a relatively closed economy with flexible exchange rates. All variables are allowed to have a contemporaneous impact on the exchange rate. The exchange and interest rates, together with the price level, are expressed in terms of quarterly averages as in the data source. This helps ensure the appropriateness of our identification scheme.

[11] Campa and Goldberg (2002) provide empirical evidence of a partial short-run pass-through from exchange-rate changes to prices in most OECD countries. Their results, however, deal with import prices. The impact on final consumer prices is generally found to weaken further.

It is in the context of the monetary policy reaction function that our identification scheme relying on contemporaneous restrictions is likely to be most controversial. First, one might argue against the use of the price level in favor of the inclusion of the inflation rate in the reaction function. However, Japan has adopted the monetary policy strategy of monetary expansion until the consumer price inflation rate is at or above zero percent. A zero-percent target inflation rate effectively corresponds to a price level target. Alternatively, one might also claim that when the inflation rate reaches negative territory, the central bank, which now want to escape the deflationary trap, has greater motivation to look first at the effects of policy on changes in the price level rather than the inflation rate. Regarding the timing assumption, Kim and Roubini (2000) and Cushman and Zha (1997) assume that the central bank lacks information about prices during the same period, implying the setting of the coefficient on a_{32} to zero. However, both these studies use monthly data. For quarterly data, the validity of this claim can be doubted. In Japan, information on the consumer price index is available 28 days after the reference month. Moreover, since monetary policy affects the economy with considerable lags, forecasts of target variables are of crucial importance in interest rate setting. As in the context of forecast targeting discussed by Svensson (1999), policy focuses on forecasts conditional on the available information and a particular future path of the instrument, implying that all the available information is taken into account. However, to emphasize the focus of the Bank of Japan on the price level instead of output fluctuations during the time frame of our study (characterized by disinflation and deflation), we set the coefficient on contemporaneous output a_{31} to zero. This could also be justified on the basis of the timing of GDP data, which is only available in the middle of the second month after each quarter. Note further that, due to the contemporaneous nature of the identification scheme, no restrictions are imposed on the lags of the effects.[12]

The identification scheme and corresponding restrictions are based on exchange rate arrangements and other considerations related to the monetary regimes of the three economies – not directly on the various theories for using the exchange rate channel to escape from a liquidity trap. Therefore, we only comment on whether the results obtained support the importance of the exchange rate channel in the recent deflationary episodes, without validating the exact predictions of the theoretical models by empirical evidence. This is due in particular to the peg of the Hong Kong dollar and the Chinese renminbi to the US currency during our examination period. Of course, a currency peg may be loosened and a more active exchange rate management provided. However, if

[12]Despite the relatively closed nature of the Japanese economy and the flexible exchange rate arrangement, an argument could be made that the BOJ follows a semi-peg for the exchange rate, whereby the exchange rate plays a significant role in the objective function of the monetary authority. In particular, the unprecedented (7% of GDP) foreign exchange interventions between January 2003 and March 2004 point to this possibility. Of course, in the case of interventions, the exchange rate is not targeted via the interest rate instrument (the BOJ policy rate was already effectively zero at the time). However, using data for an earlier period (1979-1994), Andrade and Divino (2005) argued that the BOJ aimed at stabilizing the exchange rate.

Contrasting evidence about the importance of the exchange rate in the monetary policy reaction function is provided by Jinushi et al. (2000) for the period 1975–1985. An identification scheme consistent with the objective of exchange rate stabilization leads to setting the coefficient on exchange rates (a_{34}) in the monetary policy reaction function unrestricted. The resulting system is just-identified and can be considered a robustness test for the results obtained by estimating the benchmark model for Japan.

these results are used for evaluating the feasibility of alternative monetary policy regimes, the Lucas critique may be of major importance. It is unclear that the behavior of economic agents, and thus our estimates obtained in a backward-looking scheme such as a vector autoregression, would remain robust for alternative policy arrangements.

The identification scheme for Hong Kong differs from Japan's only in the reaction function of the monetary authority (the third equation in the system below). Since October 1983, Hong Kong has enacted a currency board system with the US dollar as the anchor currency. The value of the Hong Kong currency is set at HK$ 7.8 per US$ 1. This type of arrangement, of course, represents an extreme form of a currency peg. A monetary authority in a small open economy is likely to react quickly to exchange rate shocks, which can be seen in strong interest rate movements.[13] We therefore leave unrestricted the coefficient on the exchange rate a_{34} in the monetary policy reaction function.[14] Indeed, as interest rate setting is totally endogenous in a currency board arrangement, the policy is unresponsive to local conditions in terms of real output and the price level. This supports setting of coefficients a_{31} and a_{32} in the interest rate equation to zero.[15] The model for Hong Kong is now written as:

$$\begin{bmatrix} 1 & 0 & 0 & 0 \\ a_{21} & 1 & 0 & 0 \\ 0 & 0 & 1 & a_{34} \\ a_{41} & a_{42} & a_{43} & 1 \end{bmatrix} \begin{bmatrix} u_t^y \\ u_t^p \\ u_t^i \\ u_t^{neer} \end{bmatrix} = \begin{bmatrix} b_{11} & 0 & 0 & 0 \\ 0 & b_{22} & 0 & 0 \\ 0 & 0 & b_{33} & 0 \\ 0 & 0 & 0 & b_{44} \end{bmatrix} \begin{bmatrix} \varepsilon_t^y \\ \varepsilon_t^p \\ \varepsilon_t^i \\ \varepsilon_t^{neer} \end{bmatrix} \quad (5)$$

Finally, we have the identification scheme for China. Even if the currency peg to the US dollar would, under ordinary circumstances, lead to the characterization of the Chinese monetary authority as an exchange rate targeter, capital controls might in principle allow for the pursuit of an independent monetary policy.[16] Indeed, China's capital controls include restrictions on futures trading of

[13] The degree of openness during the last three decades as measured in terms of the ratio of exports plus imports to GDP is around 200% for Hong Kong and 20% for Japan (Lin and Lee, 2002).

[14] Admittedly, the nominal effective exchange rate used in our analysis is not pegged, reflecting the influence of external forces on Hong Kong price levels. Even so, the behavior of the Hong Kong Monetary Authority well qualifies it as an exchange rate targeter, supporting our restriction.

[15] In contrast to the interest rate series for Japan, Hong Kong and Chinese interest rates are defined as the rate prevailing in the end of the period. The price level and exchange rate are period averages. While it might have been conceptually preferable to use a period average for the interest rate in the exchange rate equation, the system performed worse with such a variable in terms of residual autocorrelation for Hong Kong. Moreover, it can simply be assumed that the exchange rate is a forward-looking asset price and the end-of-period interest rate is obtained through a perfect foresight assumption. No conceptual problem arises in the monetary policy reaction function using either of the interest rate series.

[16] Officially, the exchange rate regime of China during our examination period is a managed float. In early 1994, the official and swap rates were unified at RMB 8.7 = US$ 1, with the exchange rate free to move within +- 0.25% of the previous day's reference rate (Huang and Wang, 2004). The exchange rate appreciated to RMB 8.28 = US$ 1 by October 1997, where it remained until 2005. An intense public debate about the possible undervaluation of the renminbi finally saw a revaluation of the Chinese currency by 2.1% against the US dollar in July 2005, when the PBoC announced it was abandoning the dollar peg and pegging the renminbi to a basket of currencies.

the renminbi, foreign borrowing by Chinese enterprises, portfolio investment in China by foreigners, and portfolio investment abroad by Chinese citizens (Roberts and Tyers, 2003). Since restrictions apply to capital exports as well as imports, savings have to be used inside of China. Korhonen (2004) observes that an increase in interest rates (a contractionary monetary policy shock) can have a perverse effect in such an environment by encouraging saving and thus investment, and thereby contributing to economic overheating. Applying, as for Japan, a closed-economy assumption to China, we set the contemporaneous coefficient on the exchange rate in the monetary policy reaction function (a_{34}) to zero. While the closed-economy assumption, together with the regime of flexible exchange rates, motivates the setting of the contemporaneous coefficient on exchange rates in the Japanese monetary policy reaction function to zero, an identical assumption is justified in China's case by the existence of capital controls – and despite the US-dollar peg. Further, as we are using monthly data, availability constraints regarding output data (only available with a lag) justify the setting of the coefficient on output in the monetary policy reaction function (a_{31}) to zero as we did with Japan. The overidentified model for China is then written identically to Japan's (see system (4) above).[17]

The SVAR model is estimated by maximum likelihood with respect to the matrices A and B, subject to the restrictions imposed in the structural form of the system. Numerical optimization methods are used in the form of a scoring algorithm (see Amisano and Giannini, 1997; Breitung et al., 2004). The reduced form VAR, in contrast, can be estimated by simple OLS (Japan and Hong Kong) or feasible GLS (China, due to subset restrictions).

5 Estimation Results

In presenting our estimation results, we commence with estimation of the reduced form VAR models. Results from the structural VAR analysis are reported on a country-by-country basis.

[17] Despite the restrictions in the SVAR modeling being derived from economic theory or econometric considerations, some degree of arbitrariness in the identification procedure is unavoidable. This arises because there are various ways of identifying the shocks and there is no formal statistical test to reject one correctly specified, just-identified model against another; one can merely evaluate whether the results are in line with those expected on the basis of economic theory or some conventional wisdom (e.g. expected reactions by an inflation-targeting central bank to output and price shocks). We acknowledge this drawback in the methodology and accordingly test the robustness of the results for Hong Kong and China by comparing our identification scheme to a simple recursive model (for Japan, robustness is tested by including the exchange rate in the monetary policy reaction function, as explained previously).

In the case of a recursive system, we can write $A = I_K$ and $u_t = B\varepsilon_t$. When the B matrix is restricted to be lower triangular, the first component of ε_t, ε_{1t}, can have an instantaneous impact in all the four equations, whereas the second component ε_{2t} can have an effect on all the other equations except the first, and so on. Using the same ordering of the variables as above, the exchange rate (real output) can then be considered the most (least) responsive to changes in economic conditions. Admittedly, an identification scheme obtained by recursive ordering bears a close resemblance to the contemporaneous restrictions imposed in our model for equations other than the monetary policy reaction function. However, the recursive identification scheme has been defended in the context of reduced form VAR modeling on the basis of economic considerations and has indeed been extensively used in the literature (for a model with exchange rates, see e.g. Eichenbaum and Evans, 1995).

The autoregressive order of the model is determined by the results from the misspecification tests and the need for a sufficient number of lags to satisfactorily capture the dynamic interaction between the variables. An important additional consideration is the short sample size, which can possibly lead to a weak estimation precision in a system with four endogenous variables. The models for Japan and Hong Kong are estimated using 4 lags, corresponding to one year of data, while 6 lags are included for China, where monthly data are used. Due to the high number of insignificant coefficients (common in high-order VAR models), the model for China is reduced to a subset model using a procedure where the parameter with the lowest t-value in the model is checked and possibly eliminated from the system. Such a procedure is feasible in a system where all included variables are stationary. A threshold value of 1 is specified; only coefficients with t-values above this magnitude are maintained in the model for China. A constant and trend are included as deterministic terms for all the countries. The model for Japan includes a shift dummy that obtains a value of one from 1997Q2 onward and zero before. This dummy variable corresponds to the consumption tax hike in April 1997 that induced a shift in the consumer price index. We argue it is preferable to explicitly take this tax hike into account in our short sample. As little movement in the price level is otherwise observable, this gives a rather prominent shift in the series. For Hong Kong, we include a shift dummy variable obtaining the value of 1 from 1998Q3, when instability from the Asian crisis induced a rapid decline in the inflation rate.

Misspecification and stability tests for the reduced form subset VAR models are listed in Appendix D. We perform the Portmanteau and LM-tests for residual autocorrelation, the Jarque-Bera test for non-normality and ARCH-LM tests to detect possible ARCH effects in model residuals. Neither of the tests for autocorrelation, nor the test for ARCH effects, suggest concern about model adequacy. Only for China do we find evidence of ARCH in the third equation at a 5% significance level. Since the asymptotic properties of VAR estimators are not dependent on the normality assumption, a finding of non-normality in the Jarque-Bera test is likely of minor importance for our purposes.[18] No concern about the stability of our estimated model was revealed by the CUSUM test, which is based on the cumulative sum of model residuals; the test statistic never crosses the critical lines in any of the model equations for the three economies either at 1% or 5% significance levels. Nevertheless, it is probably prudent to take such evidence at a descriptive value only for Japan and Hong Kong. Our variables are found to be integrated of order one in unit root testing and appear in

[18] In the case of Japan, we additionally perform linearity tests as proposed by Teräsvirta (1994, 2004), wherein the linearity hypothesis is tested against a smooth transition regression model. These tests are justified by the possibility of nonlinear behavior of the interest rate near the zero floor with the caveat that the validity of the tests is problematic with integrated variables.

We conduct tests on each of the four equations of the VAR model. To avoid singularity of the moment matrix, we set the deterministic terms to only appear linearly. Similarly, whenever output, prices, and the exchange rate are used as exogenous variables, they are restricted to appear linearly. Thus, we assume that it is indeed the interest rate that behaves in a nonlinear manner. The nominal interest rate was accordingly used as the transition variable from one regime to another. For the price level equation, we find weak evidence of nonlinearity when the nominal interest rate at the fourth lag was used as the transition variable. Even in this case, linearity cannot be rejected at the 1% level. We cannot reject the null hypothesis of linearity for any of the other three equations for any of the four lags of our transition variable, justifying the use of our linear model. These results are available upon request.

undifferenced form in the system for these two economies, so the validity of the CUSUM-type tests is uncertain (see Lütkepohl, 2004). Instead, we use a Chow forecast test with bootstrapped p-values. Here, stability of the estimated system is never rejected for Japan and China at a 5% significance level, while for Hong Kong smaller p-values are detected for three various break dates out of the total number of 27 tested break dates. As most of our evidence points to stable models, we proceed with the estimation of the structural systems regarding the models as reasonably adequate for our analysis.

5.1 Japan

For Japan, the structural parameter estimates of the A matrix are written as follows. To make interpretation of the contemporaneous coefficients easier, we present the negation of the A matrix (-A matrix). Thus, e.g. a positive coefficient on the price level in the monetary policy reaction function effectively corresponds to an increase in interest rates in response to a price hike.

$$
\text{-A}=\begin{bmatrix} -1 & 0 & 0 & 0 \\ \begin{matrix}-0.0033\\(0.0401)\end{matrix} & -1 & 0 & 0 \\ 0 & \begin{matrix}8.6899\\(13.4398)\end{matrix} & -1 & 0 \\ \begin{matrix}-0.7867\\(0.8477)\end{matrix} & \begin{matrix}-1.6259\\(2.9982)\end{matrix} & \begin{matrix}0.0171\\(0.0314)\end{matrix} & -1 \end{bmatrix}
$$

Asymptotic standard errors are reported in parenthesis, while the parameter estimates for the B matrix for all three economies are presented in Appendix E. A formal likelihood ratio test does not reject the overidentifying restrictions; the test statistic is $\chi^2(1) = 0.04$ with a p-value of 0.85. While the point estimates of the contemporaneous impacts perhaps hold little interest, their signs deserve discussion. Price level obtains the expected positive sign in the monetary policy reaction function (coefficient on $-a_{32}$). A price shock leads to an instantaneous exchange rate depreciation ($-a_{42}$). Similarly, a contractionary monetary policy shock leads to an expected yen appreciation ($-a_{43}$). Two counterintuitive instantaneous impacts are an exchange rate depreciation in response to an output shock ($-a_{41}$) and the negative effect on the price level of an output shock ($-a_{21}$). The finding for the exchange rate depreciation in response to an output shock is less surprising, however, considering the recent fluctuations in the yen's nominal exchange rate. Even with the slowdown in the Japanese economy and the onset of deflation, the yen appreciated notably in 1995. The rapid acceleration in economic growth in 1996 was then accompanied by exchange rate depreciation. Overall, we observe that the responses mostly obtain the expected signs, but with very limited statistical significance. More information may then be derived from a structural impulse response analysis. Taking advantage of the better small sample properties of bootstrap confidence intervals compared to other asymptotic methodologies, Hall bootstrap percentile 95% confidence intervals are used to illustrate parameter uncertainty (see Benkwitz et al., 2001). Responses up to 20 quarters ahead are considered, using 5,000 bootstrapping replications.

We now look at the responses to exchange rate and monetary policy (interest rate) shocks. The impacts of these shocks are depicted in Figure 5 below; monetary policy and exchange rate shocks appear in the left and right columns, respectively. An appreciation shock in the nominal effective exchange rate leads to a fall in real output with a statistically significant impact. Interestingly, the same shock causes a significant fall in the price level. It should be noted that the magnitude of the impact of the exchange rate shock on prices appears to be quite low. The point estimates of the impulse responses suggest that a one percent increase in the nominal effective exchange rate lowers the price level by only 0.02 percent. This maximum impact is obtained after five quarters have passed from the shock. Our estimate for the exchange rate pass-through is identical to that found by Faruqee (2004) for the euro area economy, which has the same low degree of openness as Japan. Our result matches the findings of Gagnon and Ihrig (2004) for Japan during 1981–2003.[19]

The negative impact of an exchange rate shock on the interest rate is expected in the short run; an exchange rate appreciation is contractionary from a domestic policy view and would likely be met with an interest rate lowering by the monetary authorities. The fact that this is not the long-run outcome during our estimation sample may be a result of the zero interest rate floor and limited movements in interest rates caused by the zero bound. Curiously, the yen appreciated even in the midst of a recession. Bernanke (2000) argues that such movements suggest an anticipation by speculators of an even greater deflation rate and yen appreciation in the future. The effects of an exchange rate depreciation in the "Foolproof Way" suggested by Svensson (2001, 2003) manifest themselves in both a higher volume of exports and import prices and through increasing inflation expectations in the private sector. Our results indicate that all these channels are operative in the Japanese economy – with a caveat that expectations are not captured by our system. We nevertheless obtain the result of an actual fall in the price level as a response to exchange rate appreciation.

[19] Ito et al. (2005) suggest an even lower coefficient (0.01). Neither this estimate nor Gagnon and Ihrig's (2004) is statistically significant for Japan.

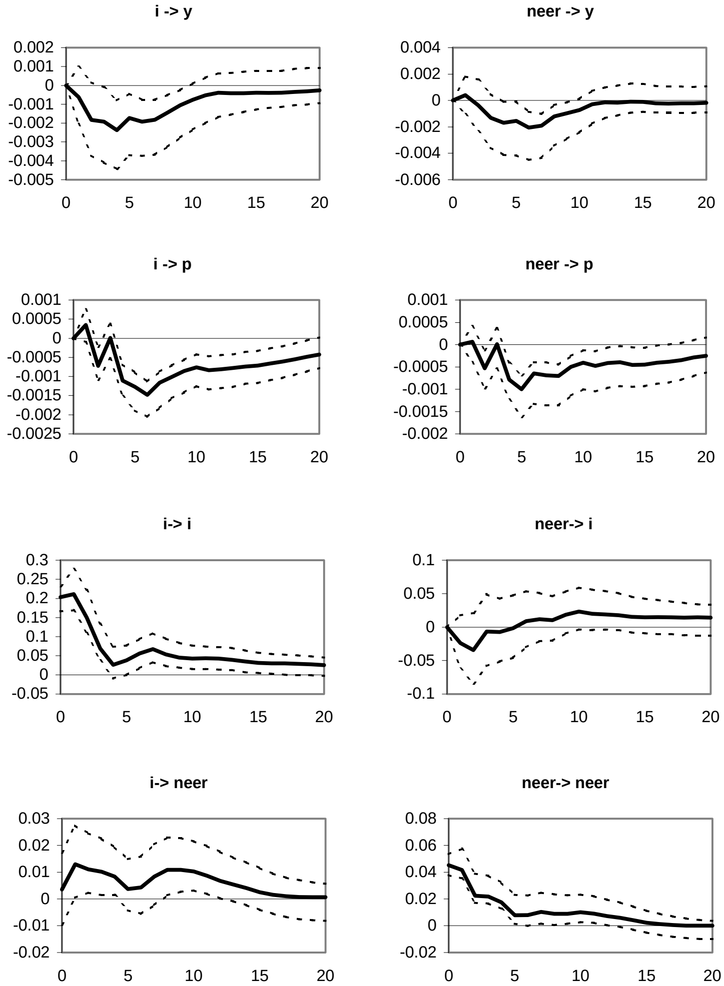

Figure 5. Impulse responses of output y, prices p, interest rate i and nominal effective exchange rate $neer$ to a structural monetary policy shock (left column) and an exchange rate shock (right column), Japan.

As predicted by Keynesian models with price-wage inertia and by liquidity models with flexible prices and wages (see Christiano and Eichenbaum, 1992, for a closed-economy setup), a contractionary monetary policy shock leads to a significant fall in real output. In accordance with an assumption of a transitory effect of a monetary contraction, output returns slowly to its pre-shock level. There is no price puzzle in our system; the price level declines significantly in response to an interest rate shock.[20] The real interest rate, which ultimately matters for aggregate demand, remains persistently positive after the period of the shock, for no price puzzle arises. The increase in the return of domestic assets

[20]This could be taken as evidence that our identified monetary policy shocks are truly exogenous in the sense suggested by Kim and Roubini (2000); we ought not witness a price level increase if the monetary contraction is not a systematic response to some other shock (such as an oil price shock or other type of inflationary pressure). Due to the model structure, a positive shock to the nominal interest rate represents a monetary contraction also defined in terms of the real rate, as the price level is not affected during the period of the shock.

leads to exchange rate appreciation, which is persistent and qualifies as delayed overshooting.[21]

In summary, both the interest rate and exchange rate channels are found to be operative in the Japanese economy during the period when interest rates declined toward the zero bound and deflation seized the economy. The potency of the interest rate channel is notable; the movements in the nominal interest rate are rather limited and close to the zero lower bound. Our findings appear to support those of Ahearne et al. (2002), who find little support for the notion that monetary policy was ineffective in the early 1990s in warding off deflationary tendencies. In a similar vein, McCallum (2003) observes that a comparison of interest rates prescribed by a Taylor rule and actual values for the overnight call rate make BOJ policy appear tight during the period 1993Q1–1998Q4. This finding of the potency of the interest rate channel supports the strategy suggested by Eggertsson and Woodford (2003) and Jung, Teranishi and Watanabe (2001) of keeping nominal interest rates at zero and committing to such a strategy for the long term (i.e. even after inflation returns). In line with the arguments of Buiter and Panigirtzoglou (2003) and Goodfriend (2000), additional stimulus can be implemented through negative interest rates by taxing currency.

Importantly, the results can be seen as weakly supportive of the strategy suggested by Svensson (2001, 2003) of using the exchange rate channel to escape from a liquidity trap. In our case, however, this is only justified on the basis of the effects of the exchange rate shocks in the system, not by the model's structure. Exchange rate shocks in our system could simply arise in the foreign exchange market without being controlled by the central bank. It is important also to keep in mind that if the exchange rate is included in the (contemporaneous) central bank reaction function, the effect of the exchange rate on prices is only borderline significant at a 5% level. But perhaps the biggest reason for caution is the low exchange rate pass-through to consumer prices suggested by the analysis. The strategy proposed by McCallum (2000, 2003), whereby the central bank follows a rule for the exchange rate, is more difficult to justify on the basis of our model (again due to the interpretation of the interest rate equation as a central bank reaction function). The nominal effective exchange rate in our model is an asset price allowed to react contemporaneously to new information in terms of output, prices, and interest rates. To the extent movements in the exchange rate reflect central bank interventions in the foreign exchange market, a McCallum-type rule would be implicitly supported by our results.[22]

[21] If exchange rate stabilization is included as an objective in the (contemporaneous) monetary policy reaction function for Japan, the signs on all previous coefficient estimates in the structural parameter matrix A remain robust. Further, the coefficient for the exchange rate ($-a_{34}$) in the interest rate equation obtains, as expected, a negative (but statistically insignificant) sign; a contractionary movement in the exchange rate is met with a fall in the interest rate. The structural impulse responses to monetary policy and exchange rate shocks under the alternative identification scheme are depicted in Appendix F. The only notable difference is the lower statistical significance of the impact on prices of an exchange rate shock at a 5% level. Otherwise, no major differences between the results yielded by the two models are observed. At a 10% significance level, a statistically significant impact is obtained during three of the 20 quarters examined. Caution therefore applies when considering manipulation of the exchange rate to escape a liquidity trap.

[22] Our empirical results regarding exchange rate shocks differ from those of Miyao (2000), who finds that an appreciation shock (defined similarly to our model in terms of the nominal effective exchange rate) leads to an increase in Japanese output, while the benchmark system does not include a price variable of any kind. Morsink and Bayoumi

5.2 Hong Kong

For Hong Kong, the structural parameter estimates for the A matrix (consistent with the approach for Japan, we use the -A matrix to facilitate interpretation) are written as follows:

$$-A= \begin{bmatrix} -1 & 0 & 0 & 0 \\ \begin{matrix} 0.0067 \\ (0.0414) \end{matrix} & -1 & 0 & 0 \\ 0 & 0 & -1 & \begin{matrix} -13.2062 \\ (23.2792) \end{matrix} \\ \begin{matrix} -0.3739 \\ (0.1438) \end{matrix} & \begin{matrix} -0.5675 \\ (0.4784) \end{matrix} & \begin{matrix} 0.0040 \\ (0.0060) \end{matrix} & -1 \end{bmatrix}$$

Asymptotic standard errors are reported in parenthesis. The likelihood ratio test does not reject the overidentifying restrictions in the case of Hong Kong either (the test statistic yields $\chi^2(1) = 0.64$, with a p-value of 0.42). As in the case of Japan, we confirm the limited significance of the instantaneous impacts in our system. The impact response of the price level to a shock in real output is positive (coefficient on $-a_{21}$). The interest rate falls in response to a shock in the nominal effective exchange rate ($-a_{34}$). Similarly, the exchange rate appreciates in response to a contractionary monetary policy shock ($-a_{43}$), and falls following a positive output shock ($-a_{41}$) and a price shock ($-a_{42}$). Thus, most instantaneous impacts have the expected signs, even if independent monetary policy is restricted by the currency board arrangement.

We next proceed to examining the dynamics in the context of impulse response analysis. The effects in the system to exchange rate and monetary policy (interest rate) shocks are depicted below in Figure 6.

(2000) also report that including the exchange rate in their VAR system often yields perverse results, e.g. a positive output response to an exchange rate appreciation. Kwon (1998) finds a depreciation shock to the yen-dollar exchange rate causes a fall in output and an increase in consumer prices. In the work by Andrade and Divino (2005), a depreciation shock to the real effective exchange rate (or the rate of deviation from its PPP value) is found to lead to lower inflation.

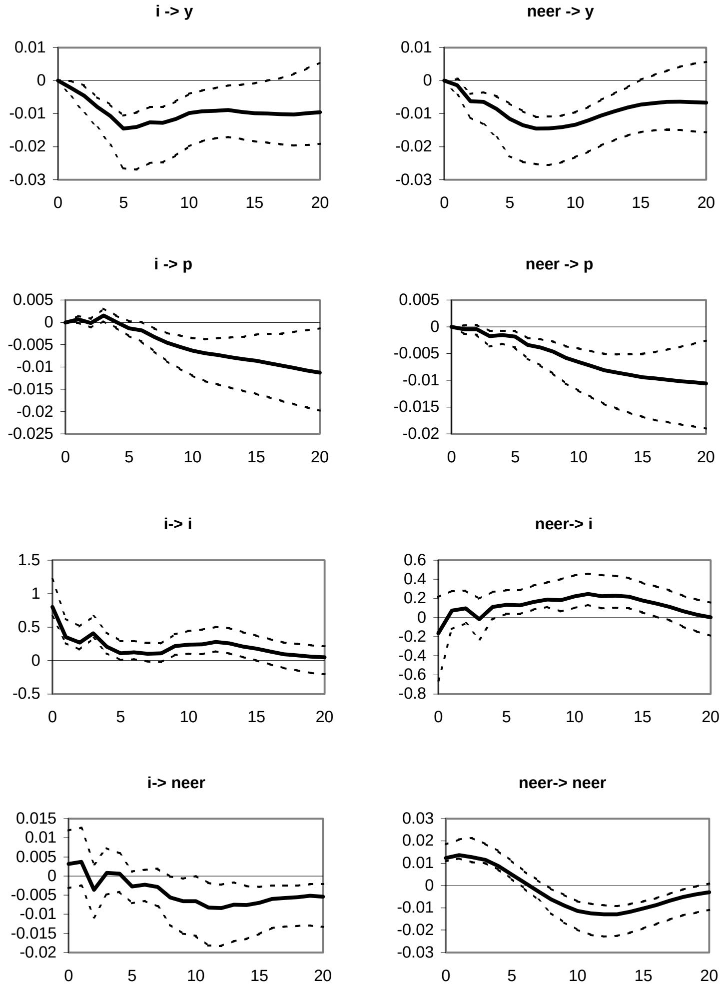

Figure 6. Impulse responses of output y, prices p, interest rate i and nominal effective exchange rate *neer* to a structural monetary policy shock (left column) and an exchange rate shock (right column), Hong Kong.

An appreciation shock in the nominal effective exchange rate leads to a fall in real output. Similar to the finding for Japan, the effect is significant, and more importantly, the exchange rate shock also leads to a significant fall in the price level. The maximum point estimate of the pass-through to consumer prices during 20 quarters amounts to 0.77, significantly higher than for the case of Japan. This high pass-through is likely the result of the extremely open nature of Hong Kong's economy. Our finding is consistent with the result of Parsley (2003), who finds a one-to-one pass-through from the nominal exchange rate to Hong Kong import prices. Since the interest rate adjusts to keep the exchange rate between the US and Hong Kong dollar constant under the currency board, it should not react to the effective exchange rate. So how do we explain why an effective exchange rate shock causes a change in interest rates? We note that, as the Hong Kong interest rate follows the US rate, the nominal effective exchange rate of the Hong Kong dollar is entirely determined outside of Hong

Kong. Apparently, this produces a counterintuitive positive impact over the long run. The effects of a contractionary monetary policy shock appear quite plausible. An interest rate shock leads to a significant fall in real output. Again, no evidence of a price puzzle is detected in our system; we see a significant decline in the price level in response to the monetary policy shock. While the increase in the return of domestic assets leads to an initial exchange rate appreciation, the exchange rate depreciates over the long run.[23]

Regarding the monetary policy shock, the fact that the responses of the price and output variables have the expected appearances clearly does not mean that the policy implied by the US dollar peg was ideal for Hong Kong during the disinflationary and deflationary era. Faced with falling prices, it is quite likely the interest rate instrument would have been wielded more strongly under an independent monetary policy. The evidence is provided in Table 1 below in the form of annual inflation and the 3-month interest rates for Japan and Hong Kong. Note, as the Japanese consumer price inflation recorded a low negative value in early 1996, the 3-month CD rate stood at 0.56%. In contrast, with an annual inflation rate of -1.03% prevailing in the Hong Kong economy in early 1999, the 3-month rate was at 6.50%. Moreover, when inflation was running at -5.29% a year later, the nominal interest rate still amounted to 5.94![24]

Time	Inflation, Japan	3-month rate, Japan	Inflation, Hong Kong	3-month rate, Hong Kong
1995M1	0.48	2.33	10.35	7.44
1996M1	-0.50	0.56	6.59	5.63
1997M1	0.51	0.53	6.29	5.56
1998M1	1.81	0.95	5.42	11.38
1999M1	0.20	0.69	-1.03	6.50
2000M1	-0.79	0.12	-5.29	5.94
2001M1	-0.40	0.48	-1.50	5.13

Table 1. Annual inflation and 3-month interest rates, Japan and Hong Kong

5.3 China

For China, the structural parameter estimates for the A matrix (consistent with the approach for the two other economies, we present the coefficients from the -A matrix to facilitate interpretation)

[23] It should be emphasized here that the exercise is somewhat counterfactual for the Hong Kong SAR. Its currency board arrangement precludes the possibility of influencing its nominal effective exchange rate to induce a positive change in the price level. Thus, we are unable to comment directly on the desirability of a strategy to depreciate the Hong Kong dollar to counteract deflationary tendencies. However, the results should be seen in the light of emphasizing the importance of internal vs. external influences on the development of Hong Kong consumer prices. Additionally, they could indirectly yield inferences about the utility of a currency board arrangement in an environment of falling prices.

[24] The contribution of the currency board to the deflationary episode has been somewhat downplayed by the Hong Kong Monetary Authority (2002). Nevertheless the HKMA conceded that about half of the fall in the CPI could be attributed to the collapse in property prices that fed directly or indirectly into the CPI, i.e. CPI components more likely to be influenced by domestic monetary conditions than, say, cyclical factors.

are written as follows:

$$
\text{-A}=\begin{bmatrix} -1 & 0 & 0 & 0 \\ \begin{matrix}-8.8201\\(4.8905)\end{matrix} & -1 & 0 & 0 \\ 0 & \begin{matrix}0.0069\\(0.0560)\end{matrix} & -1 & 0 \\ \begin{matrix}-0.0689\\(0.0938)\end{matrix} & \begin{matrix}0.0003\\(0.0019)\end{matrix} & \begin{matrix}-0.0038\\(0.0034)\end{matrix} & -1 \end{bmatrix}
$$

Again, asymptotic standard errors are in parenthesis. The likelihood ratio test for overidentifying restrictions does not suggest a rejection of our model; the test statistic is found to be $\chi^2(1) = 0.56$, with a p-value of 0.45. Similarly to the case of Hong Kong and Japan, most of the instantaneous impacts in the system are not statistically significant at conventional levels. The impact response of inflation to a shock in real output is negative (coefficient on $-a_{21}$). As expected, the interest rate is found to increase in response to a shock in the inflation rate ($-a_{32}$). All the contemporaneous responses of the exchange rate are counterintuitive in the case of China; the nominal exchange rate is found to depreciate in response to an output shock ($-a_{41}$) and a contractionary monetary policy shock ($-a_{43}$). Moreover, it is found to appreciate following a price shock ($-a_{42}$). Even if the significance of all the contemporaneous shocks is quite limited, such an outcome could well follow from China's regime of capital controls, whereby (purely external) exchange rate movements are disconnected from the domestic economy.

We depict the impulses to exchange rate and monetary policy (interest rate) shocks in the Chinese economy in Figure 7. As the underlying series appear in first differences, all the impulses have been accumulated to afford the relevant response in levels. Note, however, that the price response is the one of the annual inflation rates, so that the dynamics are not fully comparable to those of Japan and Hong Kong. Given the regulated nature of interest rates in China and their small movements, especially during the recent deflationary episode, it is quite remarkable that the responses to monetary policy shocks are consistent with conventional perceptions of such shocks, i.e. both output and the inflation rate fall. However, the impact on prices is only borderline significant at a 5% level. Moreover, the negative impact on output is no longer significant seven months after the shock. The rapid impact of a monetary policy shock on output in our system is somewhat worrisome, even if the signs of the shocks provide evidence that we have actually identified a contractionary monetary policy shock in the Chinese economy. The limited significance observed at a 5% level could simply illustrate the fact that interest rates have not held a particularly prominent status as monetary policy instruments for the PBoC. Indeed, this remains the case even after the PBoC has increasingly turned to adjusting interest rates to achieve stabilization (Zhang and Wan, 2002).

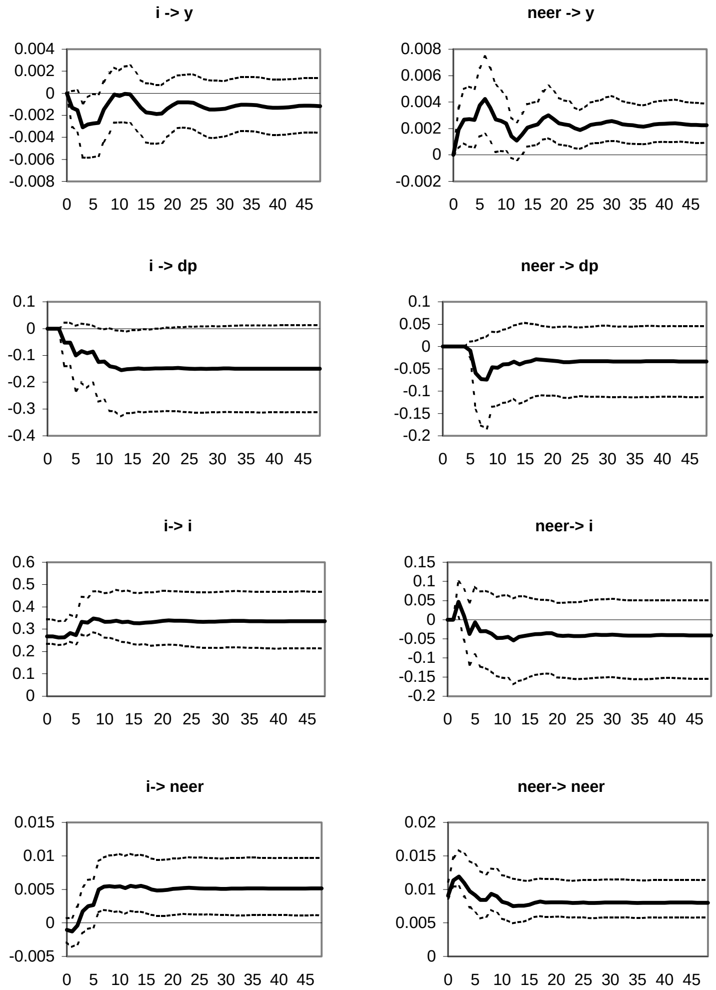

Figure 7. Impulse responses of output y, annual inflation rate Δp, interest rate i and nominal effective exchange rate *neer* to a structural monetary policy shock (left column) and an exchange rate shock (right column), China.

While an appreciation shock in the nominal effective exchange rate leads to a fall in the inflation rate, the point estimate suggests that a one percent appreciation in the nominal effective exchange rate would lead to a fall in the inflation rate by only 0.06 percentage points (obtained eight months after the shock). Further, the effect is not statistically significant. A counterintuitive finding is the effect of the exchange rate shock on output – real GDP is found to increase. This could be due to a number of factors. The contractionary impact of the renminbi appreciation (in nominal effective terms) might be offset by accommodative monetary policy; we witness a fall in the interest rate as a result of an exchange rate shock. Moreover, the shock could provide a boost for the importing sector due to lower import prices as the annual inflation rate falls. However, it seems most likely that the capital controls in the Chinese economy have to some extent insulated the domestic non-financial sector from changes in the nominal effective exchange rate. Such a finding is interesting in the context of vocal demands (particularly from the US) for Chinese renminbi appreciation. We find that such a measure could actually lead to an increase in Chinese output, even if nothing is of course suggested about the impact on the Chinese-US trade balance.[25] Finally, the impact of a contractionary monetary policy shock on the nominal exchange rate merits consideration. In the context of a theoretical model, Sun and Ma (2005) suggest that a deregulated market interest rate could work as an automatic stabilizer to alleviate part of the renminbi appreciation pressure, possibly sustaining the fixed exchange rate regime and increasing the efficiency of the banking system. This argument is supported by our empirical results to the extent that a fall in the nominal interest rate is found to lead to a statistically significant renminbi depreciation (obtained by reversing the sign of the interest rate shock in Figure 7).[26]

5.4 Forecast Error Variance Decomposition for the Three Economies

To examine the importance of different shocks on the price level (and specifically the inflation rate in China's case), we perform a forecast error variance decomposition. The procedure calculates the contribution of one variable to the forecast error variance of another variable h periods ahead (see Breitung et al., 2004). As we are primarily interested in the contributions of the various structural shocks on prices, we display only those shocks in Table 2 below. In discussing the results of this exercise, it must be kept in mind that estimation uncertainty is not tackled in the procedure.

[25] Yang and Tyers (2001), in contrast, suggest on the basis of model simulations that had the yuan been floated during the Asian financial crisis, GDP growth would have been faster.

[26] The degree of robustness for our results for Japan is investigated by considering an alternative identification scheme where the monetary authority is assumed to contemporaneously react to the nominal effective exchange rate. For the two other economies, robustness tests are performed by considering a simple recursive identification scheme. Appendix F depicts the responses to monetary policy and exchange rate shocks with alternative model structures. For Hong Kong, there is a slight change in the structural parameter estimates of the A matrix when the Cholesky decomposition is used; the nominal effective exchange rate is found to depreciate following a positive shock to output. The coefficient for output in the monetary policy reaction function is found to be positive, while that for the price level is counterintuitively negative. This likely reflects the fact that Hong Kong interest rates have been determined endogenously without considering the local economic environment in terms of output or price developments. Of course, Hong Kong business cycles may be correlated with those of the US. This would explain the positive coefficient obtained for the output variable in the reaction function. The model dynamics in the impulse response analysis for Hong Kong appear robust to an alternative identification scheme. The previously obtained structural parameter estimates for the A matrix remain robust with a Cholesky decomposition for China. Moreover, the output variable, now unrestricted, obtains the expected positive coefficient in the monetary policy reaction function. Finally, as in the case of Hong Kong, the model dynamics in the context of an impulse response analysis remain robust to the different identification scheme.

Country	Horizon	shock y	shock p (Δp for China)	shock i	shock *neer*
Japan	1	0.00	1.00	0.00	0.00
	4	0.20	0.68	0.08	0.04
	8	0.13	0.32	0.39	0.16
	12	0.25	0.22	0.38	0.15
	16	0.32	0.18	0.37	0.14
	20	0.32	0.16	0.37	0.14
Hong Kong	1	0.00	1.00	0.00	0.00
	4	0.35	0.57	0.03	0.04
	8	0.53	0.28	0.07	0.12
	12	0.43	0.16	0.19	0.22
	16	0.38	0.12	0.23	0.27
	20	0.35	0.10	0.26	0.29
China	1	0.03	0.97	0.00	0.00
	4	0.03	0.93	0.03	0.01
	8	0.03	0.93	0.03	0.01
	12	0.03	0.93	0.03	0.01
	16	0.03	0.93	0.03	0.01
	20	0.03	0.93	0.03	0.01

Table 2. Proportions of forecast error in the consumer price index (Japan and Hong Kong) and the annual inflation rate (China).

Regarding Japan, we see that monetary policy shocks, measured by shocks to the interest rate, dominate the shocks in the consumer price index over the long run. The proportion of inflationary shocks in the forecast error variance of consumer prices is quite high in the short run, but declines strongly over time. In contrast, the importance of exchange rate shocks remains relatively constant and is highest two years after the shock.

For Hong Kong, the results resemble those of Japan. Both monetary policy and inflationary shocks are found to contribute quite significantly to movements in consumer prices. Like in Japan, the importance of the latter decreases over time. Notably, however, the importance of exchange rate shocks is high in the long run – in fact, twice as high in magnitude as that in Japan after five years.

Table 2 suggests that the Chinese inflation rate is mainly determined by inflationary shocks in both the short and long run. Conventional monetary policy shocks and shocks to the exchange rate are of minor importance and vary little over time. This is in stark contrast to their role detected in the two other economies (monetary policy shocks, especially in Japan's case, and exchange rate shocks particularly for Hong Kong).

What might explain the relatively high importance of the exchange rate shocks in Japan in the short run (i.e. up to two years after the shock)? Campa and Goldberg (2002) remark that the exchange rate pass-through to import prices is considerably larger for Japan – a country with a large share of raw materials and energy in its imports – than for other major industrial countries. In the longer run, however, we find the importance of exchange rate shocks to be higher in Hong Kong than in Japan. This is in line with the limited insulation properties of fixed exchange rate regimes against shocks to foreign inflation. Also, while we witnessed a negative and significant impact on prices of an exchange rate appreciation shock for Hong Kong, the impact for Japan was only of a weak statistical significance when an alternative identification scheme was used. The use of the exchange rate channel in Japan should thus be seen in the light of increasing inflation expectations and lowering the real ex ante interest rate, as in the theoretical zero bound literature, and not merely as a way to stimulate exports and lower import prices. This is even more the case, as the extent of the pass-through from the nominal effective exchange rate to consumer prices (measured in the context of impulse response analysis) is found to be significantly lower in Japan than in Hong Kong.

Our results for Hong Kong are broadly in line with the findings of Genberg (2003), i.e. we confirm the relatively high importance of the nominal effective exchange rate (determined completely outside Hong Kong) for the price level in the long run, and find domestic monetary conditions play a relevant role, as evidenced by the importance of monetary policy shocks for price movements during the disinflationary and deflationary era.[27] Of course, due to increased economic integration, price convergence between Hong Kong and the mainland could be of importance for our results. Ha and Fan (2004) found evidence that up to one fifth of Hong Kong's deflation could be attributed to such structural adjustment. When these shocks are represented by "own" inflationary shocks in our forecast error variance decomposition, their share in the system is rather small in the long run, amounting to only 10% after five years.

For China, which also has an exchange rate peg, our results suggest exchange rate shocks have very limited importance in the determination of consumer prices. This finding is consistent with the low share of fuel imports and a large share of manufactures in China's import composition,[28] price regulation, and, perhaps most importantly, the existence of capital controls that in principle would allow for an independent monetary policy. Our results are somewhat in contrast with Ha et al. (2003), who find that the value of the renminbi and world prices are important determinants of long-term price movements in China, with the former explanatory variable obtaining a prominent role during the era of low inflation and deflation. Yet, even if the nominal effective exchange rate does not appear to be a significant factor in our analysis, world prices could still well be of importance. Indeed, these may actually be captured by the inflationary shocks in our system. These shocks have (perhaps surprisingly) prominent importance during our estimation sample; their share consistently accounts for over 90% of overall shocks. The limited use of a conventional interest rate

[27] The Hong Kong Monetary Authority (2002) has argued that there is no clear evidence of real interest rates having a strong influence on current expenditure on goods and services in Hong Kong, partly since durable goods are imported to the economy. Our findings of the importance of monetary policy shocks (defined in terms of nominal interest rates) could be seen to contradict this claim.

[28] In China, the share of manufactures of total imports amounted to 80.2% in 2002, while fuel imports stood at only 6.5%.

instrument in the Chinese monetary policy framework is reflected in the results from our forecast error variance decomposition, where monetary policy shocks account for only a tiny fraction of the overall movements in consumer prices. Our findings, especially the contrasting results for the two economies with fixed exchange rates, illustrate that the institutional and economic differences ranging far beyond the exchange rate arrangement may have major importance.

6 Conclusions

Our paper set out to examine the role of the interest rate and exchange rate channels during the recent deflation episodes in three closely interlinked economies: Japan, Hong Kong, and China. We estimated an open-economy structural vector autoregressive (SVAR) model with contemporaneous restrictions for three economies, which all use different exchange rate regimes and monetary policy arrangements. For Japan and Hong Kong, an appreciation shock in the nominal effective exchange rate leads to a statistically significant fall in the price level, with a considerably higher impact on the Hong Kong economy. Similarly, the impact of the interest rate shock on prices is rather strong in both economies. In contrast, neither exchange rate nor interest rate shocks significantly influence price development in China. The limited importance of these shocks is hardly surprising as Chinese monetary policy has been predominantly implemented through administrative measures rather than market-determined interest rates. Similarly, Chinese capital controls may have largely insulated the economy from foreign shocks originating in the exchange rate.

Our results suggest that appreciation of the nominal effective exchange rate could represent a powerful external deflationary factor for an economy, or alternatively, that depreciation of the currency could provide a way to escape from the liquidity trap in situations where use of the conventional interest rate channel is limited. The latter approach is in line with suggestions in the theoretical literature, most prominently the "Foolproof Way" recommended by Svensson (2001, 2003). With the pursuit of currency pegs in Hong Kong SAR and China, our results concerning the importance of the nominal effective exchange rate channel should be taken as an inference about the role of external factors for price movements in these economies. Notably, the importance of monetary policy shocks for consumer price movements in Hong Kong and Japan was relatively high during our estimation period when short-term interest rates were close to or at the zero bound. In such an environment (assuming measures to induce negative interest rates are not considered), the interest rate channel could still be operative through an aggressive monetary easing before inflation becomes negative. Alternatively, the central bank could use the expectations channel by committing to keeping interest rates low for a considerable period of time as suggested by Eggertsson and Woodford (2003), and Jung, Teranishi and Watanabe (2001).

If multiple economies with flexible exchange rates or adjustable pegs find themselves in a liquidity trap, they cannot simultaneously depreciate against one another. This was recognized by Svensson (2001, 2003). However, in the situation where Hong Kong and China were using fixed dollar pegs, Japan was in a position to depreciate its currency against the dollar. Admittedly, a more active exchange rate management in Japan might trigger accusations of "beggar-thy-neighbor" policies, especially where the required exchange rate adjustment is necessarily substantial.

Our results yield no information about the transmission of monetary shocks (for interest or exchange rates) from one economy to another; the nominal effective exchange rate takes into account multiple trading partners as an aggregate measure. The extension of the model to allow for a more detailed international transmission of interest rate, exchange rate, and price shocks is left for future study.

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APPENDIX A
DATA SOURCES

The following variables were obtained from the OECD Main Economic Indicators Database: Real GDP (Japan), consumer price index (Japan, China) and the 3-month interest rate on the certificate of deposit (Japan).

The nominal effective exchange rate for China is from the IFS Database (series IFS.M.924.NECZF.)

Thomson Datastream was used as the source for the following series: the 3-month interbank rate (Hong Kong, series HKINTER3, original source: Hong Kong Monetary Authority); nominal GDP (China, series CHGDP...A, original source National Bureau of Statistics of China); average repo rate (China, series CHYREPOA, original source People's Bank of China); annual inflation rate (China, series CHCONPR%F); nominal effective (trade-weighted) exchange rate (Japan, series JPQ...NEUE; Hong Kong, series HKQ...NECE); consumer price index (Hong Kong, series HKQ64...F).

The GDP series for Hong Kong was obtained from Hong Kong Statistics, Census and Statistics Department.

The author made seasonal adjustments to the Chinese nominal GDP series, the Hong Kong consumer price index, and Hong Kong real GDP using a Census X-11 procedure. The consumer price index and real GDP for Japan were seasonally adjusted at the data source. Chinese nominal GDP was linearly interpolated to monthly observations and deflated by the consumer price index to obtain an estimate of real output at a monthly frequency. The Chinese consumer price index was calculated by assuming linear growth in consumer prices for the year 1993, and setting a value of 100 for January 1993. Monthly observations were subsequently calculated using the monthly year-on-year change in consumer prices.

FIGURES OF SERIES USED IN ESTIMATION

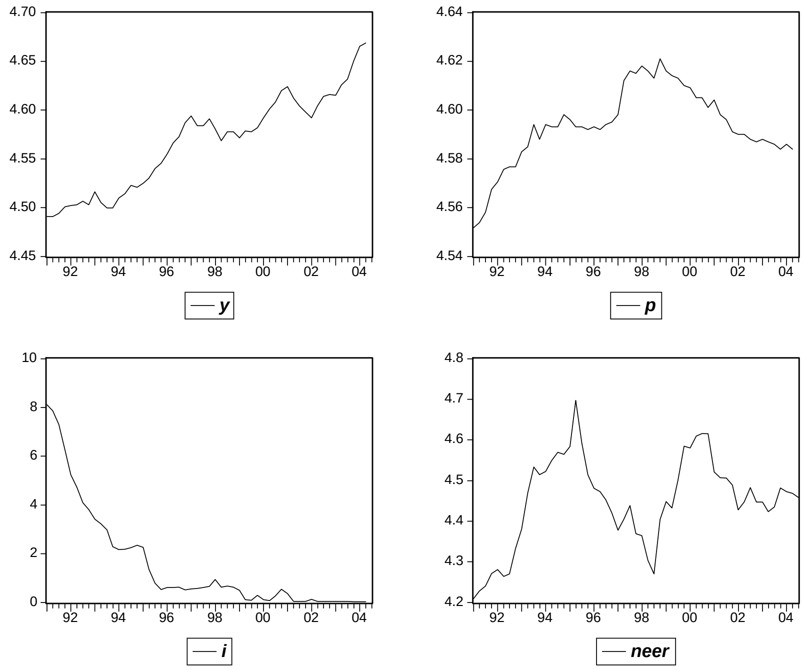


Figure: Series, Japan.

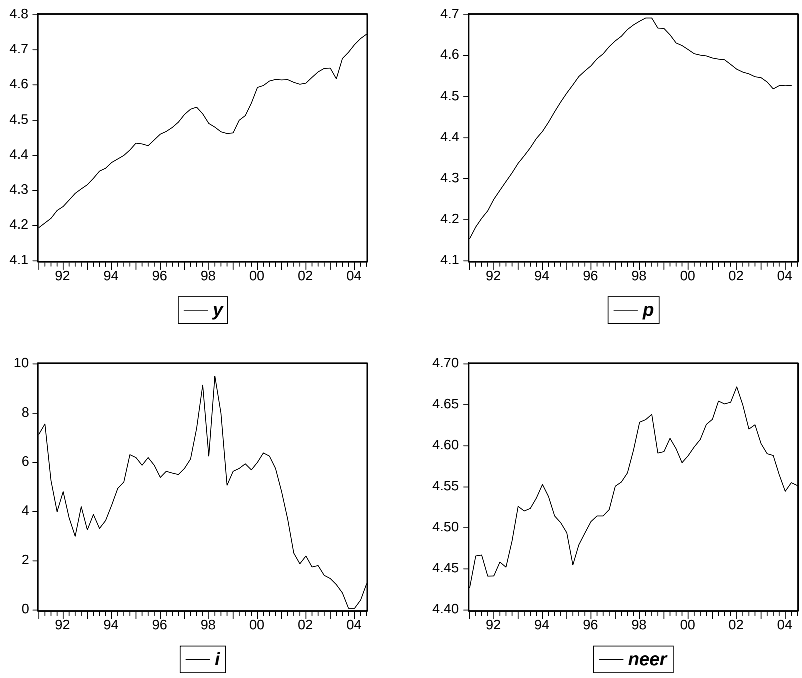


Figure: Series, Hong Kong.

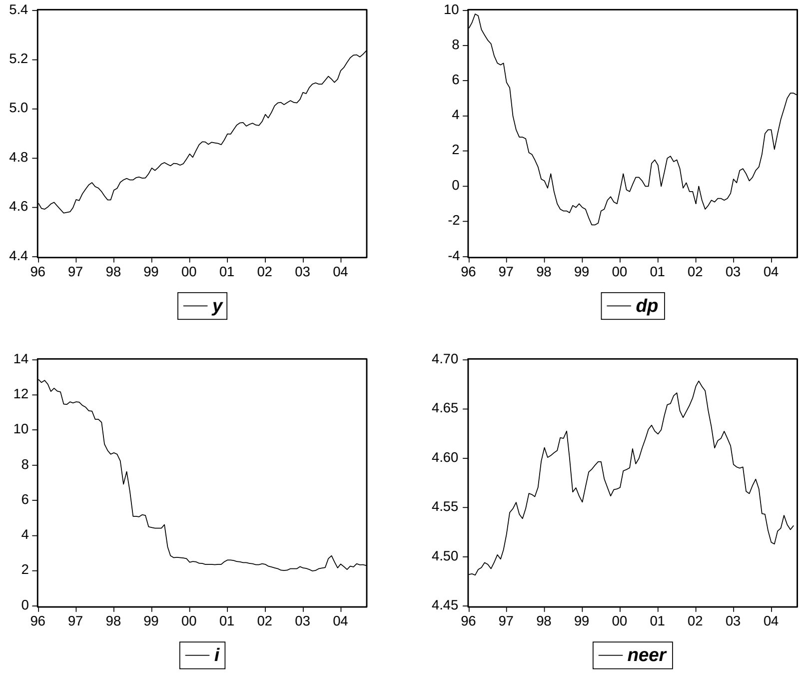


Figure: Series, China.

APPENDIX B

UNIT ROOT TESTS

Augmented Dickey-Fuller Test, Hong Kong

Series	Det. term	Lagged differences	Test stat.
$\Delta^2 y$	c	5 (AIC,HQ,SC)	-4.00***
Δy	c	0 (AIC,HQ,SC)	-5.32***
y	c, t	3 (AIC)	-3.25*
y	c, t	2 (HQ,SC)	-3.20
$\Delta^2 p$	$c, S98Q3$	3 (AIC, HQ, SC)	-1.69
Δp	$c, S98Q3$	0 (AIC,HQ,SC)	-3.00**
p	c, t	3 (AIC, HQ, SC)	-2.53
$\Delta^2 neer$	c	7 (AIC)	-3.55***
$\Delta^2 neer$	c	5 (HQ,SC)	-3.11**
$\Delta neer$	c	0 (AIC,HQ,SC)	-5.77***
$neer$	c, t	3 (AIC)	-2.31
$neer$	c, t	1 (HQ)	-1.39
$neer$	c, t	0 (SC)	-1.10
$\Delta^2 i$	c	6 (AIC)	-2.48
$\Delta^2 i$	c	3(HQ,SC)	-4.07***
Δi	c	0 (AIC,HQ,SC)	-9.04***
i	c, t	0 (AIC,HQ,SC)	-1.83

* indicates significance at 10% level, ** at 5% and *** at 1% level.

Order specification criteria in parenthesis: AIC=Akaike, HQ=Hannan-Quinn, SC=Schwarz criteria.

Price series tested with unit root test with structural break (Lanne et al., 2002)

Prefix S denotes date of shift dummy

c and t denote constant and trend as deterministic terms, respectively.

All series except interest rates in logarithms.

1991Q1-2004Q3, maximum number of lags set at 10.

Augmented Dickey-Fuller Test, Japan

Series	Det. term	Lagged differences	Test stat.
$\Delta^2 y$	c	5 (AIC,HQ,SC)	-3.22**
Δy	c	6 (AIC,HQ)	-2.97**
Δy	c	0 (SC)	-5.80***
y	c, t	6 (AIC,HQ)	-3.34*
y	c, t	1 (SC)	-2.43
$\Delta^2 p$	c	8 (AIC)	-2.16
$\Delta^2 p$	c	0 (HQ,SC)	-4.34***
Δp	c	2 (AIC)	-2.49
Δp	c	0 (HQ,SC)	-7.09***
Δp	$c, I97Q2$	2 (AIC)	-2.98**
Δp	$c, I97Q2$	0 (HQ,SC)	-7.32***
p	c, t	0 (AIC,HQ,SC)	-1.66
p	$c, t, S97Q2$	1 (AIC,HQ)	-0.53
p	$c, t, S97Q2$	0 (SC)	-0.85
$\Delta^2 neer$	c	5 (AIC)	-2.63*
$\Delta^2 neer$	c	3 (HQ)	-2.99**
$\Delta^2 neer$	c	2 (SC)	-2.47
$\Delta neer$	c	0 (AIC,HQ,SC)	-3.20**
$neer$	c, t	3 (AIC)	-2.66
$neer$	c, t	0 (HQ,SC)	-2.06
$\Delta^2 i$	c	4 (AIC,HQ,SC)	-2.57*
Δi	c	0 (AIC,HQ,SC)	-3.59***
Δi	c	0 (AIC,HQ,SC)	-8.41***
i	c, t	1 (AIC,HQ,SC)	-4.36***

* indicates significance at 10% level, ** at 5% and *** at 1% level.

Order specification criteria in parenthesis: AIC=Akaike, HQ=Hannan-Quinn, SC=Schwarz criteria.

Price series additionally tested with unit root test with structural break (Lanne et al., 2002) where the prefixes S and I denote the dates of the shift and impulse dummy variables, respectively.

c and t denote constant and trend as deterministic terms, respectively.

All series except interest rates in logarithms.

1991Q1-2004Q2, maximum number of lags set at 10.

Augmented Dickey-Fuller Test, China

Series	Det. term	Lagged differences	Test stat.
$\Delta^2 y$	c	10 (AIC,HQ,SC)	-2.65*
Δy	c	5 (AIC,HQ,SC)	-7.84***
y	c, t	10 (AIC,HQ)	-0.11
y	c, t	6 (SC)	-2.11
$\Delta^3 p$	c	8 (AIC,HQ)	-0.88
$\Delta^3 p$	c	2 (SC)	-3.29**
$\Delta^2 p$	c	6 (AIC,HQ)	-1.61
$\Delta^2 p$	c	0 (SC)	-8.71***
Δp	c, t	0 (HQ,SC)	-0.87
$\Delta^2 neer$	c	8 (AIC,HQ)	-2.01
$\Delta^2 neer$	c	4 (SC)	-3.68***
$\Delta neer$	c	0 (AIC,HQ,SC)	-7.34***
$neer$	c, t	1 (AIC,HQ,SC)	-1.46
$\Delta^2 i$	c	6 (AIC,HQ,SC)	-1.93
Δi	c	0 (AIC,HQ,SC)	-9.44***
i	c, t	0 (AIC,HQ,SC)	-0.51

* indicates significance at 10% level, ** at 5% and *** at 1% level.

Order specification criteria in parenthesis: AIC=Akaike, HQ=Hannan-Quinn, SC=Schwarz criteria.

c and t denote constant and trend as deterministic terms, respectively.

All series except interest rates and the inflation rate in logarithms.

Δp denotes the annual (year-on-year) inflation rate.

1996M1-2004M8, maximum number of lags set at 10.

APPENDIX C. Saikkonen-Lütkepohl Cointegration Test, China

Series	Deterministic term	no. of lags	Cointegration rank	test statistic
$y, i, p, neer$	c, t	7 (AIC)	0	35.46
			1	16.99
			2	9.90
			3	0.81
$y, i, p, neer$	c, t	1 (HQ, SC)	0	33.40
			1	15.92
			2	5.83
			3	0.04
y, i, p	c, t	7 (AIC)	0	23.58
			1	5.14
			2	0.44
y, i, p	c, t	1 (HQ,SC)	0	24.07
			1	5.04
			2	0.02
$y, i, neer$	c, t	7 (AIC)	0	12.70
			1	5.54
			2	0.78
$y, i, neer$	c, t	2 (HQ)	0	16.30
			1	4.96
			2	0.39
$y, i, neer$	c, t	1 (SC)	0	13.49
			1	2.68
			2	1.16
$y, p, neer$	c, t	2 (AIC,HQ)	0	23.99
			1	5.84
			2	0.08
$y, p, neer$	c, t	1 (SC)	0	19.67
			1	8.36
			2	0.06
$i, p, neer$	c, t	2 (AIC)	0	19.81
			1	5.97
			2	0.05
$i, p, neer$	c, t	1 (HQ,SC)	0	16.72
			1	5.25
			2	0.15

* indicates significance at 10%, ** at 5% and *** at 1% level.

c and t denote constant and trend as deterministic terms, respectively.

Order specification criteria in parenthesis: AIC=Akaike, HQ=Hannan-Quinn, SC=Schwarz criteria.

APPENDIX D

MISSPECIFICATION AND STABILITY TESTS

Japan

Q_{16}	207.92 [0.20]
FLM_5, FLM_4, FLM_1	1.28 [0.21], 1.53 [0.06], 1.04 [0.42]
$LJB(s_3^2), LJB(s_4^2)$	1.08 [0.90] 3.81 [0.43]
$ARCH_{LM}(16)$(eqs. 1, 2, 3, 4)	18.17 [0.31] 18.18 [0.31] 16.12 [0.44] 21.02 [0.18]

Hong Kong

Q_{16}	179.97 [0.72]
FLM_5, FLM_4, FLM_1	1.42 [0.12], 1.23 [0.22], 1.54 [0.11]
$LJB(s_3^2), LJB(s_4^2)$	3.59 [0.46] 16.10 [0.00]
$ARCH_{LM}(16)$(eqs. 1, 2, 3, 4)	11.36 [0.79] 5.93 [0.99] 17.19 [0.37] 20.14 [0.21]

China

Q_{16}	203.11 [0.77]
LM_5, LM_4, LM_1	56.97 [0.98], 41.61 [0.99], 8.61 [0.93]
$LJB(s_3^2), LJB(s_4^2)$	21.94 [0.00] 17.66 [0.00]
$ARCH_{LM}(16)$(eqs. 1, 2, 3, 4)	18.11 [0.32] 7.33 [0.97] 42.74 [0.00] 11.24 [0.79]

Note: p-values in brackets.

Q denotes the Portmanteau test statistic for autocorrelation.

(F)LM is the Lagrange multiplier type (F) test statistic for autocorrelation.

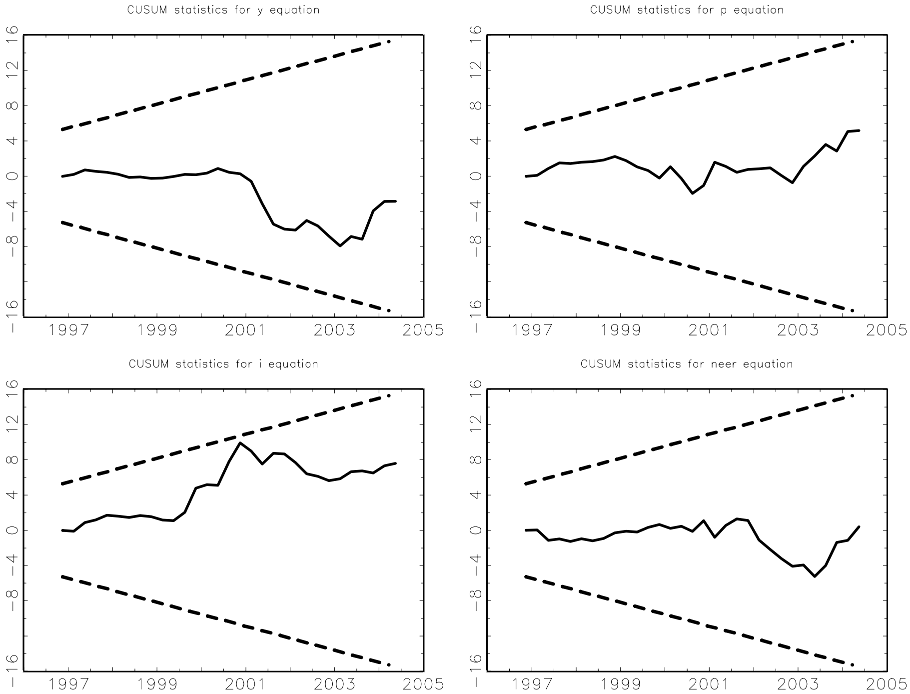
LJB is the Lomnicki-Jarque-Bera joint test for nonnormality; for skewness only (s_3^2) and kurtosis only (s_4^2), as in Lütkepohl (1991).

ARCH-LM is a Lagrange multiplier test for autoregressive conditional heteroskedasticity.

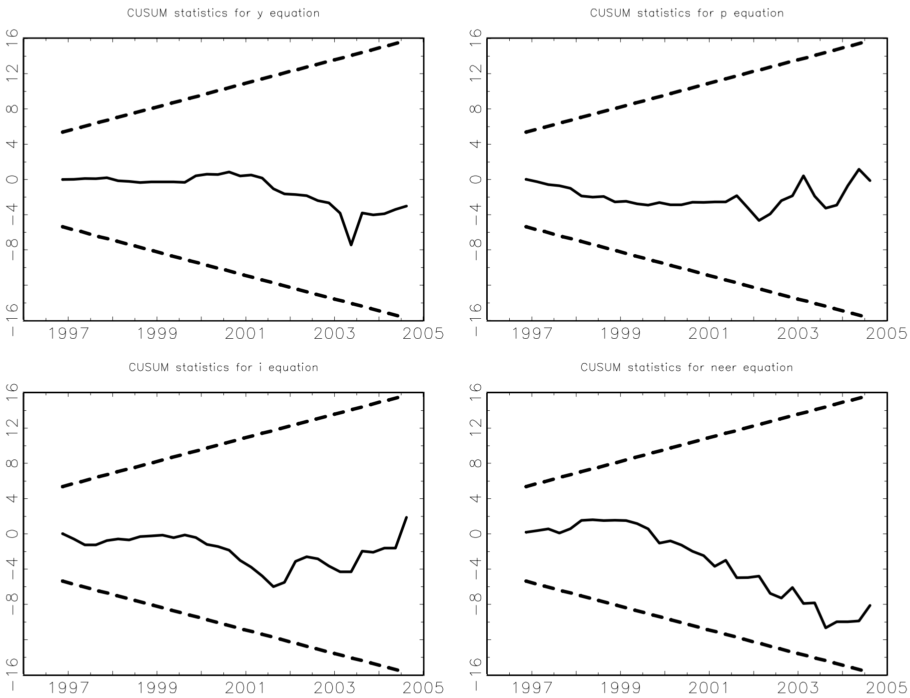
16 lags used for the Portmanteau and ARCH-LM tests, 5, 4 and 1 lags for the LM test.

STABILITY TESTS

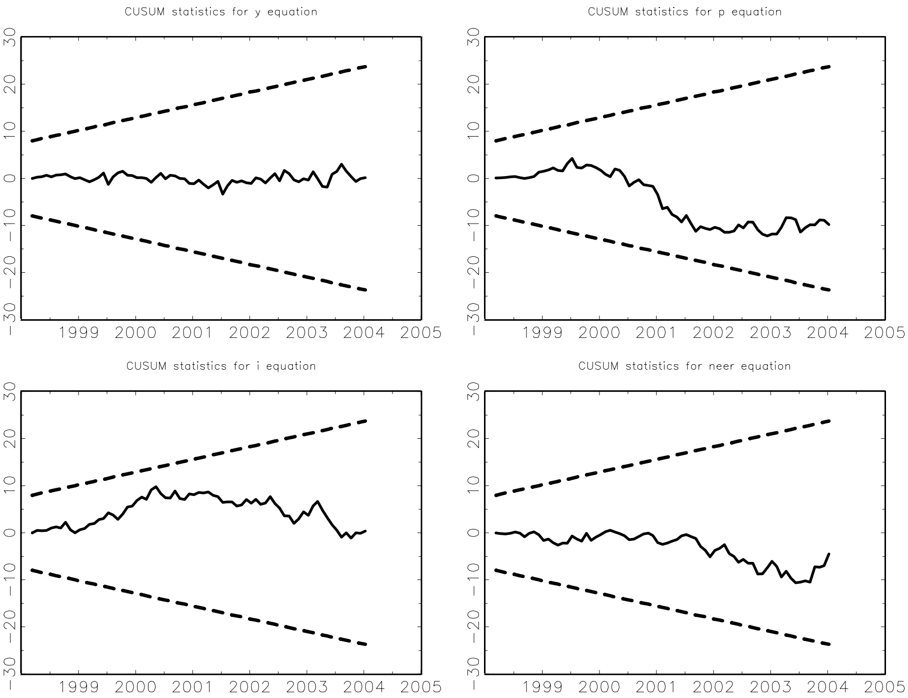
CUSUM TESTS, 5% SIGNIFICANCE LEVEL



Japan



Hong Kong


CUSUM statistics for y equation
CUSUM statistics for p equation
CUSUM statistics for i equation
CUSUM statistics for neer equation
30
20
10
0
−10
−20
−30
1999 2000 2001 2002 2003 2004 2005

China

	Japan		Hong Kong		China					
break date	test stat	p-val	test stat	p-val	break date	test stat	p-val	break date	test stat	p-val
1997 Q4	1.8687	0.439	5.2598	0.003	1999 M2	0.0464	0.726	2001 M11	0.5702	0.902
1998 Q1	1.2104	0.61	1.5989	0.117	1999 M3	0.1044	0.703	2001 M12	0.576	0.902
1998 Q2	1.4267	0.425	1.3963	0.21	1999 M4	0.1449	0.763	2002 M1	0.6028	0.819
1998 Q3	0.9053	0.846	1.4545	0.154	1999 M5	0.1849	0.783	2002 M2	0.6209	0.82
1998 Q4	0.9784	0.789	1.8553	0.157	1999 M6	0.1915	0.941	2002 M3	0.6021	0.871
1999 Q1	1.0508	0.689	1.2867	0.371	1999 M7	0.2215	0.931	2002 M4	0.6071	0.854
1999 Q2	1.1511	0.547	1.4894	0.204	1999 M8	0.2729	0.834	2002 M5	0.6208	0.839
1999 Q3	1.24	0.405	1.6191	0.14	1999 M9	0.3073	0.819	2002 M6	0.6177	0.864
1999 Q4	1.1827	0.461	1.635	0.093	1999 M10	0.3456	0.731	2002 M7	0.6491	0.782
2000 Q1	0.9362	0.737	1.2175	0.37	1999 M11	0.3771	0.679	2002 M8	0.6648	0.739
1991 Q2	0.927	0.753	1.1169	0.446	1999 M12	0.4128	0.633	2002 M9	0.6866	0.673
1991 Q3	1.0323	0.598	1.0988	0.519	2000 M1	0.4465	0.565	2002 M10	0.7195	0.601
1991 Q4	1.0298	0.591	1.2157	0.383	2000 M2	0.4612	0.579	2002 M11	0.7512	0.49
2001 Q1	1.0797	0.521	1.2264	0.341	2000 M3	0.4093	0.863	2002 M12	0.7639	0.497
1992 Q2	1.0168	0.595	1.3365	0.221	2000 M4	0.3913	0.956	2003 M1	0.8011	0.377
1992 Q3	0.7987	0.827	1.4376	0.159	2000 M5	0.4132	0.93	2003 M2	0.7918	0.412
1992 Q4	0.5815	0.974	1.3345	0.216	2000 M6	0.4007	0.968	2003 M3	0.8209	0.405
2002 Q1	0.6414	0.954	1.4087	0.19	2000 M7	0.3945	0.984	2003 M4	0.7868	0.446
2002 Q2	0.5388	0.976	1.3708	0.21	2000 M8	0.4174	0.967	2003 M5	0.8506	0.308
2002 Q3	0.5024	0.99	1.5344	0.126	2000 M9	0.4336	0.96	2003 M6	0.7888	0.489
2002 Q4	0.5061	0.972	1.7066	0.056	2000 M10	0.4551	0.941	2003 M7	0.8337	0.358
2003 Q1	0.4917	0.974	1.9987	0.015	2000 M11	0.4842	0.899	2003 M8	0.8589	0.306
2003 Q2	0.4965	0.947	2.1111	0.019	2000 M12	0.4694	0.951	2003 M9	0.9168	0.234
2003 Q3	0.5702	0.904	1.7379	0.073	2001 M1	0.4934	0.942	2003 M10	0.9398	0.217
2003 Q4	0.7176	0.737	0.8673	0.647	2001 M2	0.5019	0.926	2003 M11	0.9054	0.28
2004 Q1	0.3074	0.964	1.1452	0.358	2001 M3	0.4832	0.969	2003 M12	0.8695	0.344
2004 Q2			0.9578	0.51	2001 M4	0.502	0.947	2004 M1	0.9066	0.302
					2001 M5	0.5144	0.934	2004 M2	0.8612	0.37
					2001 M6	0.5258	0.948	2004 M3	0.8078	0.455
					2001 M7	0.557	0.876	2004 M4	0.8105	0.429
					2001 M8	0.5775	0.861	2004 M5	0.9142	0.356
					2001 M9	0.5702	0.88	2004 M6	0.5622	0.752
					2001 M10	0.5428	0.933	2004 M7	0.6368	0.62

Chow forecast test statistics for Japan, Hong Kong and China.
Bootstrapped *p*-values based on 1,000 replications

APPENDIX E

STRUCTURAL PARAMETER ESTIMATES FOR THE B MATRIX

(asymptotic standard errors in parenthesis)

$$\begin{bmatrix} \begin{matrix}0.0075\\(0.0008)\end{matrix} & 0 & 0 & 0 \\ 0 & \begin{matrix}0.0021\\(0.0002)\end{matrix} & 0 & 0 \\ 0 & 0 & \begin{matrix}0.2034\\(0.0203)\end{matrix} & 0 \\ 0 & 0 & 0 & \begin{matrix}0.0452\\(0.0045)\end{matrix} \end{bmatrix}$$

Japan

$$\begin{bmatrix} \begin{matrix}0.0129\\(0.0013)\end{matrix} & 0 & 0 & 0 \\ 0 & \begin{matrix}0.0038\\(0.0004)\end{matrix} & 0 & 0 \\ 0 & 0 & \begin{matrix}0.8435\\(0.1115)\end{matrix} & 0 \\ 0 & 0 & 0 & \begin{matrix}0.0130\\(0.0016)\end{matrix} \end{bmatrix}$$

Hong Kong

$$\begin{bmatrix} \begin{matrix}0.0099\\(0.0007)\end{matrix} & 0 & 0 & 0 \\ 0 & \begin{matrix}0.4788\\(0.0344)\end{matrix} & 0 & 0 \\ 0 & 0 & \begin{matrix}0.2683\\(0.0193)\end{matrix} & 0 \\ 0 & 0 & 0 & \begin{matrix}0.0090\\(0.0006)\end{matrix} \end{bmatrix}$$

China

APPENDIX F

STRUCTURAL IMPULSE RESPONSE ANALYSIS, ALTERNATIVE IDENTIFICATION SCHEMES

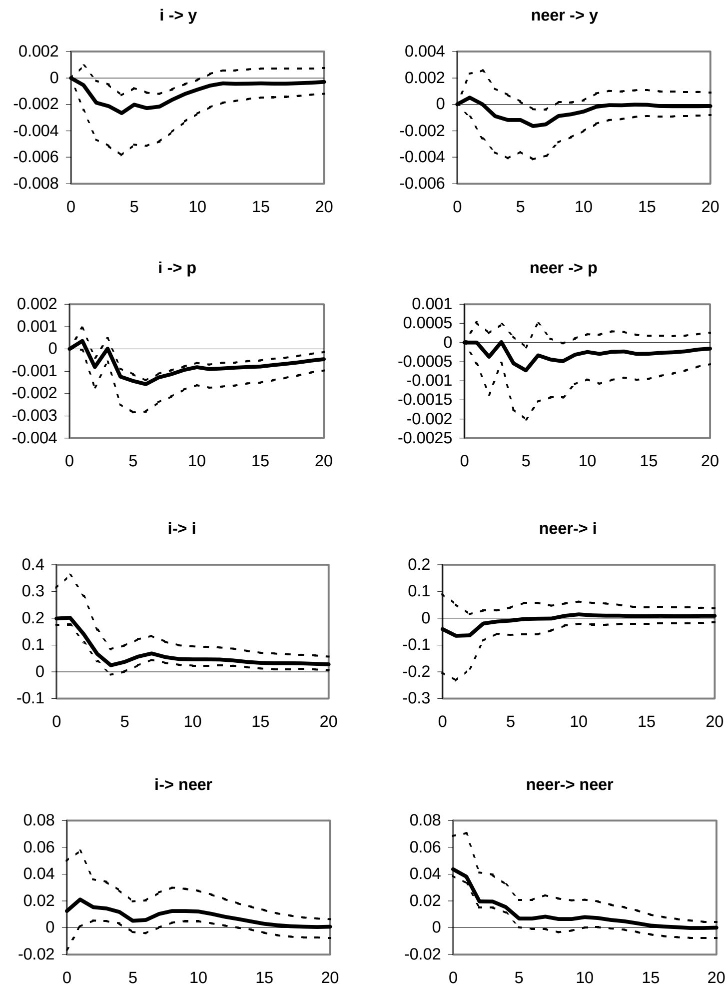


Figure: Impulse responses of output y, prices p, interest rate i and nominal effective exchange rate $neer$ to a structural monetary policy shock (left column) and an exchange rate shock (right column); Japan, exchange rate targeting identification scheme.

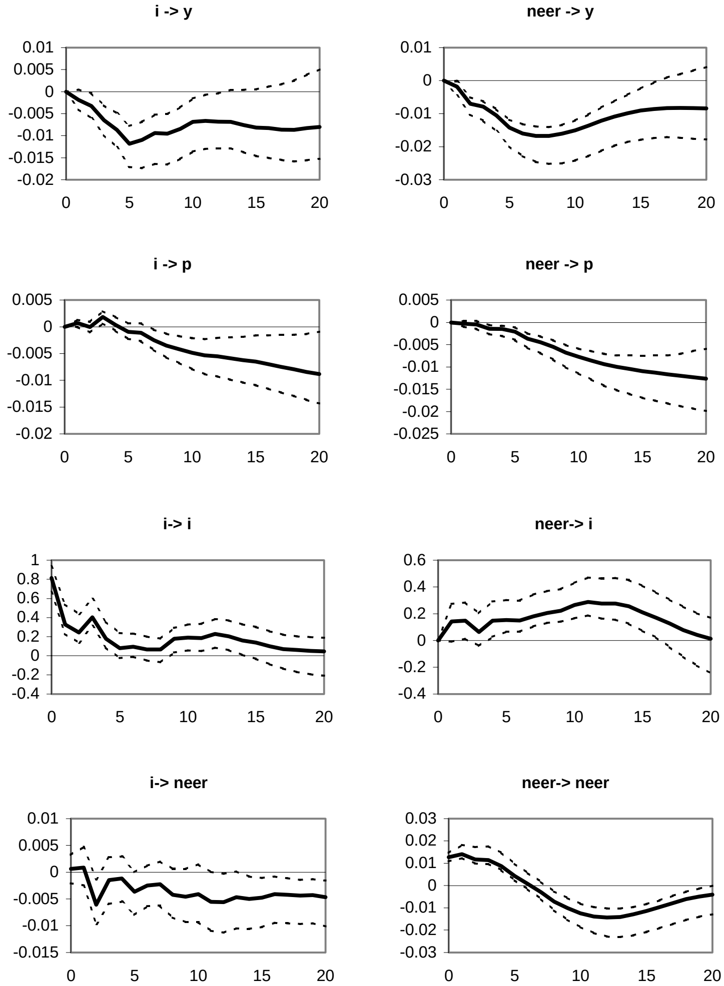


Figure: Impulse responses of output y, prices p, interest rate i and nominal effective exchange rate $neer$ to a structural monetary policy shock (left column) and an exchange rate shock (right column); Hong Kong, Cholesky decomposition.

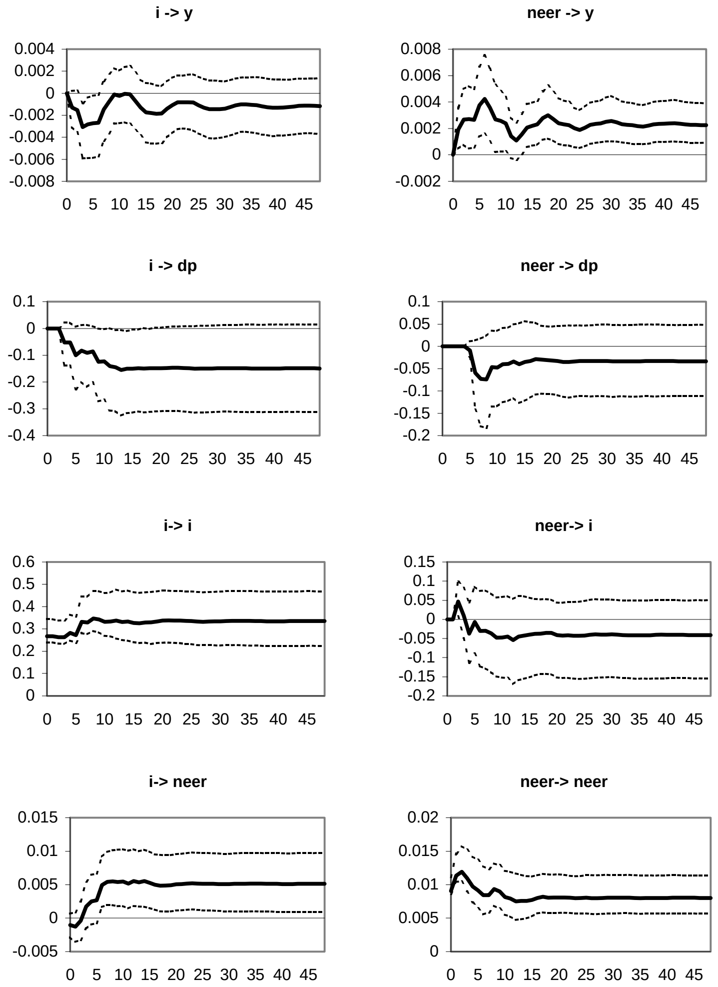


Figure: Impulse responses of output y, annual inflation rate Δp, interest rate i and nominal effective exchange rate $neer$ to a structural monetary policy shock (left column) and an exchange rate shock (right column); China, Cholesky decomposition.

2004

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