EUROPEAN UNIVERSITY INSTITUTE
Department of Economics

Essays on Empirical Macroeconomics

Aaron Mehrotra

Thesis submitted for assessment with a view to obtaining the degree of Doctor of the European University Institute

Florence, February 2006

EUROPEAN UNIVERSITY INSTITUTE
Department of Economics

Essays on Empirical Macroeconomics

Aaron Mehrotra

The Thesis Committee consists of:

Mike Artis, supervisor
Tam Bayoumi, IMF
Stefan Gerlach, University of Basel
Helmut Lütkepohl

Contents

European University Institute

Department of Economics

Ph.D. Dissertation

Essays on Empirical Macroeconomics

Aaron Mehrotra

Thesis committee:
Michael Artis
Tamim Bayoumi
Stefan Gerlach
Helmut Lütkepohl

Acknowledgements

My interest in postgraduate studies led me to the European University Institute, located in the lovely environment of Florence and San Domenico di Fiesole, in September 2001. During these years, the EUI has provided an inspiring and stimulating atmosphere to pursue research. I am greatly indebted to my supervisors Professor Michael Artis and Professor Helmut Lütkepohl for their enduring guidance and support. My supervisors have greatly inspired me to conduct research in the fields of monetary economics and applied time series econometrics. It is my sincere hope that this thesis reflects some of my enthusiasm for these topics. I also wish to express my gratitude to Professor Pertti Haaparanta at the Helsinki School of Economics who encouraged me to pursue PhD studies.

I would like to thank the co-author of the fourth thesis essay, Tuomas Peltonen, for the enjoyable cooperation. Working together has always been a pleasure.

Financial support from the Academy of Finland is gratefully acknowledged.

Finally, I owe my gratitude to my family and friends. I would like to express my warmest thanks to my parents for their constant and unfailing support and care. And to Leonor, whose encouragement and friendship have been so important during my PhD.

This thesis is dedicated to my parents, with love.

Chapter 1

Introduction

This thesis consists of four essays in empirical macroeconomics. The first three essays examine the conduct of monetary policy during a disinflationary and deflationary era, with the policy interest rates close to or at the zero lower bound. While the Japanese economic environment of the 1990s and the first years of the 21st century make the country an obvious focus for empirical study, the third essay extends the analysis to Hong Kong and China - both economies have recently experienced deflation. In contrast, the last chapter focuses on the fiscal policy aspects of the European Union's (EU) most recent enlargement. In this introduction, we elaborate on the motivation underlying the four essays and provide a brief overview of our findings, commencing with the conduct of monetary policy in a disinflationary environment and concluding with the fiscal policy and development aspects of EU enlargement.

The successful disinflationary process in industrialized countries has given life to phenomena that were previously thought of as being primarily of historical or mere theoretical interest, including deflation and the zero interest rate floor as a binding constraint on monetary policy. Nowhere has this been more apparent than in Japan, where the bursting of the asset price bubble in the early 1990s has given way to economic stagnation and falling consumer prices for most of the time since the mid-1990s. At the time of writing, despite some encouraging signs of economic recovery, deflation seems rather entrenched, even with the policy rates of the Bank of Japan effectively at zero since 1999. Not surprisingly, this experience has inspired a number of theoretical and empirical studies concerning the mechanisms, if any, through which monetary policy could be stimulative under such extraordinary circumstances.

Rather than focusing on unconventional policy measures – although some are discussed – our interest is centred on research questions that are arguably more fundamental in nature. In particular, our work empirically examines whether traditional perceptions of the effects of monetary policy hold when inflation rates are low or negative, and the policy interest rate is close to or at the zero lower bound. It is possible that the zero floor emerges as a territory where the behaviour of the macroeconomic variables of interest is unchanged, with the only difficulty (though still a major one) arising from the infeasibility of further lowering the interest rate. Alternatively, traditional economic relationships could fail to hold at the lower bound due to significant changes in the behaviour of economic agents. We consider this to be of fundamental importance, as some of the theoretical literature concerned with the conduct of policy at the zero bound fails to consider the possibility that monetary policy actions may no longer be transmitted to the real economy. This is crucial for an optimal commitment policy to keep interest rates low well into the future even when the deflationary impulse has dissipated (e.g. Eggertsson and Woodford, 2003). It is additionally of importance for a price level targeting strategy that largely relies on the real *ex ante* interest rate to provide the stimulative impact (e.g. Svensson, 2003). But even a currency tax that allows for any negative nominal interest rate (e.g. Buiter and Panigirtzoglou, 2003) would not provide any macroeconomic stimulus if the interest rate channel is blocked.

In accordance with the considerations outlined above, our research agenda includes an investigation of the potency of the interest rate channel in the Japanese economy through the disinflationary and deflationary period. Theoretically, such a research question could be disregarded outright – no theory suggests that the interest rate channel (perhaps modelled as an IS equation) weakens when inflation rates fall. Yet macroeconomic conditions in the Japanese economy may well have proven otherwise, particularly due to problems in the financial sector which saw an accumulation in bad loans and a declining creditworthiness of borrowers. Indeed, some studies point to breaks in the transmission mechanism in the mid-1990s, while surprisingly few studies examine monetary transmission well into the deflationary period. We intend to fill this gap in the literature. In addition, our research agenda includes an estimation of a broad money demand relationship in the Japanese economy during the disinflationary and deflationary era. Again, our interest focuses on a canonical relationship in monetary policy analysis, for we are interested in the existence of a stable money demand relationship in an economy that has slid from positive to negative inflation rates. However, rather than proposing a strategy of monetary targeting for the Bank of Japan, we discuss the information content of broad money and the dynamics between money, price indicators and income – this requires a stable relationship between the variables of interest. Finally, for a central bank close to or at the zero bound, the exchange rate channel may provide a rare opportunity to affect economic activity and the price level. Accordingly, we investigate the importance of both the interest and exchange rate channels during recent deflationary episodes in Japan, Hong Kong and China.

In the following, we briefly summarize the findings from the three essays concerned with the conduct of monetary policy during deflation. As our chosen methodology, we mainly utilize vector autoregressions (VARs) in the analysis. Since reduced form VARs primarily serve as tools to summarize the dynamic properties of the data, they are atheoretical in nature. However, the use of structural VARs, both with short and long-run restrictions, allows for theoretical considerations to be implemented in the analysis. Moreover, a VAR approach could be defended on the grounds that no strict consensus about the exact effects of monetary policy in the economy prevails. Probably the greater disadvantage of this methodology arises from the fact that a major part of the dynamic analysis focuses on shocks rather than systematic aspects of policy. We acknowledge this conceptual problem with the chosen approach.

The first essay of this thesis investigates whether inflation or price level targeting would be appropriate policy regimes for Japan. The feasibility of either of the two approaches critically hinges on two main prerequisites: the interest rate channel must remain potent during a period of disinflation and deflation, and the central bank must possess an instrument to influence the future price level. In addition to examining the fulfilment of these conditions, this paper can be seen as a simple approach to identifying monetary policy shocks in the Japanese economy, as a limited number of studies extend the analysis long into the deflationary period. Using evidence from both vector autoregressions and structural form IS equations, we find that the interest rate channel remains potent even with falling inflation rates. Of course, using an atheoretical approach and a backward-looking framework such as a vector autoregression, the Lucas-critique may be of importance in discussing alternative monetary policy rules for the Bank of Japan. However, as parameter constancy in our system seems to be fulfilled even with the onset of deflation, we believe that the Lucas-critique would not pose a problem to the methodology chosen. With nominal interest rates at zero, monetary expansion may be achieved through a commitment strategy whereby rates are kept low well into the future. An additional stimulus could, however, be implemented by introducing a tax on currency, since such a strategy would allow any desired negative policy interest rate to be created. As short and long-term real *ex ante* interest rates are found to be significant for the determination of output in the structural IS equations, price level targeting obtains some support in our analysis – expectations play a crucial role in the successful implementation of such a strategy.

In the second essay, we are interested in the demand for broadly defined liquidity during the Japanese disinflationary and deflationary period. Three motivations are prominent. Firstly, an investigation of this very broad measure of liquidity could be considered interesting in its own right, for we are aware of only one study that has examined this aggregate in a vector error correction framework (Bank of Japan, 2003). Secondly, only limited evidence of broad money demand extending far into the deflationary era exists. Thirdly, the current interest in broad money largely stems from considerations of its information value, especially with regard to destabilizing asset price developments that have undoubtedly been of importance in the Japanese context. Using a vector error correction framework, we find that a stable demand relationship for broadly defined liquidity can be established in the Japanese economy even through the period of financial instability and the onset of deflation. While we find equity prices to be highly important in our system, there is little evidence of significant currency substitution effects. In contrast to most of the previous literature on Japanese money demand, we implement an impulse response analysis to examine the dynamic relationships between the variables, considering two different identification schemes. Interestingly, bearing in mind the persistence of deflation in the Japanese economy, impulse response analysis suggests causality from broadly defined liquidity to both consumer and equity prices.

The final chapter concerned with the conduct of monetary policy during deflation takes into account external aspects by considering the importance of the exchange rate channel, in addition to the interest rate channel, in the recent deflationary episodes in Japan, Hong Kong and China. While all these three economies have recently experienced deflation, they are characterized by strikingly different monetary policy regimes and varying degrees of openness to the rest of the world. Estimating structural vector autoregressions with contemporaneous restrictions, we find that exchange rate shocks have had a statistically significant impact on prices in both Japan and Hong Kong during the disinflationary era, but the quantitative impact is markedly higher in the latter economy. Our findings provide evidence of the importance of external influences in these economies, and could alternatively be seen as providing weak support to suggestions to using exchange rate depreciation in order to escape from the liquidity trap (see e.g. Svensson, 2001; McCallum, 2000; Coenen and Wieland, 2003). A certain degree of counterfactuality in the analysis for Hong Kong is unavoidable, as independent monetary policy with an interest rate instrument is restricted due to the currency board arrangement, and the exchange rate cannot be used as a monetary policy tool to induce changes in the domestic price level. However, the analysis could indirectly give rise to an inference about the suitability of the currency board during a deflationary period. It is indeed plausible that independent monetary policy may have responded more strongly to falling prices. For China, neither interest rate nor exchange rate shocks are found to be of importance in determining consumer prices.

Very little of our analysis is critical of the monetary policy implemented by the Bank of Japan. The Japanese policymakers found themselves in largely uncharted territory with nominal interest rates at the zero bound, and the slide toward deflation was to a significant extent unexpected by both the domestic and international observers (Ahearne *et al.*, 2002). But whereas the picture emerging from our analysis perhaps seems unexpectedly encouraging, it is difficult to induce a further stimulative impact in the economy once the zero bound is hit. In other words, while we find that basic relationships between the variables of interest appear largely unaltered by deflation and the looming or binding zero bound, a macroeconomic policy to escape from the deflationary trap may well be difficult to implement. This could be due to political reasons (as in the case of introducing a tax on currency) or because the required stimulus may be very big (as in the case of yen depreciation). Similarly, while broad money seems to possess desirable causal properties with respect to consumer and equity prices, and the relationship between money, income and prices appears to be stable, inducing an increase in broad money is difficult especially in the case of instabilities in the money multiplier. Pre-emptive policymaking may then still provide the best alternative when (forecasts of) inflation rates appear unexpectedly low and scope for monetary easing remains. This is not to suggest that monetary policy is ineffective – for we contend that this is definitely not the case – it just emphasizes the argument that the required stimulus at the zero bound may be sizeable and politically difficult to implement.

The fourth essay moves the discussion from monetary policy during a deflationary era to the fiscal policy aspects of EU enlargement. In particular, we are interested in whether the fiscal austerity required by the Maastricht criteria and the Stability and Growth Pact would be harmful for the socio-economic catching-up process of the new Member States. To our knowledge, no previous work about the impacts of fiscal policy on socio-economic development for the new Member States exists. Most of the literature on EU enlargement has been concerned with the choice of exchange rate and monetary policy regimes, the extent of real convergence, and business cycle analysis (see e.g. Begg *et al.*, 2003). However, as many of the new Member States are still undergoing a transition process from command to market economies, with some of them falling quite far below the average EU income levels, socio-economic progress could well be of importance for these economies. If government spending and investment are efficient and beneficial for development, fulfilling the fiscal criteria may be detrimental to the new Member States, as many of them are currently running high government deficits and notable consolidation measures would be required. Alternatively, fiscal austerity could be beneficial for socio-economic development, economic growth and stability and thus ultimately for welfare.

In our study, the level of socio-economic development is evaluated by constructing a socio-economic development index (SEDI). This measure consists of various socio-economic indicators that are largely affected by public policies. Our index suggests that the cohesion countries Greece, Ireland, Portugal and Spain were approximately at the same level of socio-economic development in the 1980s, when most of them joined the European Union, as the new Member States in 1999. This allows us to use data for the cohesion countries to evaluate the relationship between socio-economic development and fiscal policy, and to draw inferences for the new Member States.

Our results from instrumental variables regressions show that fiscal consolidation would be beneficial to socio-economic development in the medium term. In line with previous literature about the effects of fiscal consolidation on economic output (see e.g. Alesina and Ardagna, 1998), we find that fiscal retrenchment, including a lower level of public debt, would be advantageous to development. Finally, we evaluate how long it would take for the new Member States to achieve the EU benchmark levels in terms of the socio-economic development indicator, assuming the average speed of development of the cohesion countries during 1980-1999. The times vary from 8.5 years (Slovenia) to 24 years (Romania). A policy implication of our results is to maintain the Stability and Growth Pact or an equivalent intergovernmental fiscal rule to curb public spending and debt.

Bibliography

[1] Ahearne, A., Gagnon, J., Haltimaier, J., Kamin, S., 2002. Preventing Deflation: Lessons from Japan's Experience in the 1990s. Board of Governors of the Federal Reserve System, International Finance Discussion Paper no. 729.

[2] Alesina, A., Ardagna, S., 1998. Tales of Fiscal Adjustment. Economic Policy 27, 487-546.

[3] Bank of Japan, 2003. The Role of the Money Stock in Conducting Monetary Policy. Bank of Japan Quarterly Bulletin May 2003, 1-52.

[4] Begg, D., Eichengreen, B., Halpern, L., von Hagen, J., 2003. Sustainable Regimes of Capital Movements in Accession Countries. CEPR Policy Paper no. 10.

[5] Buiter, W. H., Panigirtzoglou, N., 2003. Overcoming the Zero Bound on Nominal Interest Rates with Negative Interest on Currency: Gesell's Solution. The Economic Journal 113, 723-746.

[6] Coenen, G., Wieland, V., 2003. The Zero-Interest-Rate Bound and the Role of the Exchange Rate for Monetary Policy in Japan. Journal of Monetary Economics 50 (5), 1071-1101.

[7] Eggertsson, G. B., Woodford, M., 2003. The Zero Bound on Interest Rates and Optimal Monetary Policy. Brookings Papers on Economic Activity 2003 (1), 139-211.

[8] McCallum, B. T., 2000. Theoretical Analysis Regarding a Zero Lower Bound on Nominal Interest Rates. Journal of Money, Credit, and Banking 32 (November, Part 2), 870-904.

[9] Svensson, L. E. O., 2001. The Zero Bound in an Open Economy: A Foolproof Way of Escaping from a Liquidity Trap. Monetary and Economic Studies 19, 277-312.

[10] Svensson, L. E. O., 2003. Escaping from a Liquidity Trap and Deflation: The Foolproof Way and Others. Journal of Economic Perspectives 17 (4), 145-166.

Chapter 2

The Case for Price Level or Inflation Targeting - What Happened to Monetary Policy Effectiveness during the Japanese Disinflation?

2.1 Introduction

The protracted economic slump in Japan since the early 1990s has aroused the attention of economists worldwide and has led to numerous policy proposals. This so-called "lost decade", having lasted much longer than a simple cyclical downturn, has brought to life economic phenomena that have proved a veritable challenge for policy makers. Interestingly, even though inflation has almost continuously been negative since 1998 and the zero interest rate policy was first initiated in 1999, the Bank of Japan (BOJ) has been unwilling to adopt either an inflation or a price level targeting strategy. Admittedly, the current policy of quantitative easing bears some resemblance to price level targeting, as it is to be in place until CPI inflation stays at or above zero percent for a few months - a zero percent inflation rate effectively corresponds to a price level target. Even so, the absence of an explicit price level or inflation targeting strategy may seem surprising given the potential benefits of either of these approaches and the widespread adoption of inflation targeting by central banks worldwide.

The aim of our paper is to examine whether price level or inflation targeting would be appropriate policy choices for Japan. The feasibility of either strategy hinges on two main prerequisites: the interest rate must remain a potential tool for monetary policy during periods of disinflation and deflation, and the central bank must possess an instrument to actually be able to influence the future price level.[1] On the basis of results from vector autoregressions (VARs), accompanied by estimates from structural form IS equations, we find that both conditions are satisfied in the case of Japan. However, monetary transmission to economic activity and prices appears to be remarkably slow. A straightforward policy proposal to create additional stimulus is the implementation of a tax on currency, effectively allowing for any negative policy interest rate. An expansion of broad money is found to be causal for increases in the price level and real output. Finally, results from the structural equations provide evidence for the significance of the real *ex ante* interest rate and output expectations in the determination of economic activity in Japan, affording support for a price level targeting strategy where expectations play a crucial role.

Even though numerous studies have addressed Japan's deflation problem, covering fiscal, monetary and structural policy aspects, our approach is novel for a number of reasons. Firstly, by conducting estimations through end-2003, we are able to capture a significant number of years when deflation prevailed in the Japanese economy. If there were significant structural breaks in the monetary transmission mechanism as the economy moved from positive inflation rates to deflation, our estimation sample would be likely to capture these. This is crucial, since if the monetary transmission mechanism had broken down in the economy, neither inflation nor price level targeting would be feasible policy alternatives for the Bank of Japan. Indeed, a number of studies have pointed to possible breaks in monetary transmission in this economy, discussed in more detail in the following section. By providing evidence on both structural VAR estimates with contemporaneous restrictions and structural IS equations, we are able to evaluate whether monetary policy was in principle effective *ex post* as deflation took hold and additionally, whether future prices and interest rates *ex ante* are of importance in the determination of current output. The importance of the real *ex ante* interest rate is prominently emphasized in theoretical literature about price level targeting. Secondly, even if some of our policy proposals are not novel in the sense that they have been previously suggested in the literature, they are derived from our empirical results in a straightforward fashion. This contrasts

[1] Of course, any change in the monetary policy strategy of a central bank is ultimately a political decision. Ito (2004) provided interesting political economy evidence on why the BOJ has been unwilling to adopt an inflation targeting strategy.

with studies where unorthodox monetary policy measures are proposed in a purely theoretical context without any link to empirical estimates regarding the monetary transmission mechanism. Finally, and no less importantly, we feel that there is scope in the existing literature for identifying monetary policy shocks in Japan in a simple yet meaningful fashion, especially for the period when interest rates have declined toward the zero bound and deflation has taken hold of the economy.

Recent signs of recovery in the Japanese economy have increased hopes that the long period of deflation is nearing its end. Even so, the possibility of hitting the zero bound and the threat of deflation have recently caused serious concern in other big economies as well. After the Federal Funds rate was lowered to 1 percent in the US in June 2003, concerns arose about the constraint a lower bound on interest rates could place on monetary policy. Moreover, while the European Central Bank (ECB) has publicly seemed to downplay the possibility of negative inflation rates in the euro area, it did indeed argue that the clarification in May 2003 of its inflation target of "under but close to 2 percent" was aimed at creating a sufficient safety margin against deflation (ECB, 2003).

This paper is structured as follows. In the next section, we discuss some of the previous literature concerning the slowdown of the Japanese economy and its monetary transmission mechanism, together with theoretical considerations that are prominent to the analysis. This is followed by a presentation of the methodological approach of the study. In the fourth section, the empirical analysis in the form of vector autoregressions is performed and policy implications derived from the results are discussed. Furthermore, results from structural IS equations are investigated and analysed. The final section concludes.

2.2 Previous Literature and Theoretical Considerations

This section is divided into two main parts. Firstly, we discuss some previous work that has examined the Japanese slowdown and the monetary transmission mechanism in that economy. Due to the large number of studies concerned with Japan's deflation problem and in line with our chosen methodology, we focus primarily on research that has applied vector autoregressions to examine Japanese monetary policy. As we will contend, there still remains scope for a simple framework to investigate the effects of monetary policy shocks in the Japanese economy, especially during the deflationary era. Secondly, we mention some topics of interest in the theoretical literature that concern the conduct of policy near or at the zero interest rate floor. Other studies of relevance to our work are mentioned in later parts of the text, as appropriate.

2.2.1 VAR Analysis of Japanese Monetary Policy

Studies employing vector autoregressions to investigate Japanese monetary policy have focused on various issues, such as the slowdown in the 1990s, a general identification scheme to examine monetary policy shocks, and special features of the deflationary period, including an investigation of structural breaks in the monetary transmission mechanism. The study by Bayoumi (2001) belongs to the first category. The author examined the growth slowdown of the 1990s in Japan, testing for the significance of different explanations for the slump. These included the absence of a bold fiscal policy, the limitations of monetary policy, falling domestic asset prices and the disruption of financial intermediation. According to the study, all the above explanations could have some validity; however, banking system problems seemed to be the main reason for the weakness in growth. Due to a sample period finishing in the first quarter of 1998 that the author had at his disposal, the study did not include many years when deflation prevailed in the Japanese economy. In similar vein, Morsink and Bayoumi (2000) found that banks play a crucial role in transmitting monetary shocks to economic activity. These authors argued that policy measures to strengthen banks would be a prerequisite to enhancing monetary policy effectiveness in Japan.

Regarding the identification of Japanese monetary policy shocks, Sims (1992) examined many countries - Japan included - using a recursive model where shocks to short-term interest rates were identified as those of exogenous monetary policy. Similarly, the analysis of Kim (1999) was conducted in a multicountry framework for the G-7 economies, using common identifying assumptions across countries. While the restrictions imposed in the paper were of a contemporaneous nature, they differed from a simple recursive form. Shioji (2000) extended the nonrecursive system examined by Kim (1999) by including some additional variables, such as Bank of Japan loans and high-powered money. In contrast, Jang and Ogaki (2003) used a structural vector error correction model and focused on the effects of Japanese monetary policy shocks on exchange rates.

The paper by Kasa and Popper (1997) studied the objectives and operating procedures of the BOJ during 1975-1994. The authors found that the BOJ weighted variation both in the call money rate and in nonborrowed reserves, with the emphasis on the interest rate increasing over time. Miayo (2002) used a recursive VAR model to study the effects of monetary policy on aggregate activity and found that monetary policy shocks had a persistent effect on real output especially during the boom-and-bust economy of the 1980s. In another paper, the author found a break in the reduced form VAR system in the mid-1990s. More specifically, the persistent effect of a monetary policy shock on real output disappeared in the subsample of the 1990s (Miayo, 2000). This was also supported by stability testing that yielded a break date of 1995. Intuitive explanations for this finding included the appreciation of the yen, very low levels of short-term interest rates and problems in the banking system at that time.

Kimura *et al.* (2002) took into account the regime change to quantitative easing and the possible non-linearity of money demand at low interest rates, and accordingly estimated a VAR with time-varying parameters. While the positive effect of monetary base on inflation still existed in the mid-1980s, it was found to have disappeared by 2002. Similarly to Jang and Ogaki (2003), Kimura *et al.* (2002) and Miyao (2000), who examined dynamics within a vector autoregressive model with a monetary aggregate, Arai and Hoshi (2004) focused on the long-run relation between broad money M2+CDs, prices and output. Even if a cointegrating relation between broad money and real GDP was found to exist under the low interest rate regime, a break was found in the long-run relationship in the mid-1990s.

As is clear from the previous discussion, some of the VAR evidence suggests the existence of a break in the monetary transmission mechanism. Such arguments have also been made in alternative estimation frameworks. Fujiki and Shiratsuka (2002) and Okina and Shiratsuka (2003) examined the duration effect of the zero interest rate policy, derived from the central bank's commitment to maintaining interest rates low for a considerable time in the future, by estimating forward rate curves with Japanese money market data. The authors found that the policy had indeed stabilized market expectations of the future course of short-term interest rates, bringing longer-term interest rates down. However, it was argued that as the monetary policy transmission mechanism was not functioning, the effects of the policy were not felt in other parts of the economy, but this claim was not examined further.

If the monetary transmission mechanism has weakened considerably in Japan, suggestions to implement a price level or an inflation targeting strategy would greatly lose their ground. Therefore, analysis of the potency of the interest rate channel is of special interest for our purposes. But even more generally, we consider that there is room for a simple identification scheme to investigate the impacts of Japanese monetary policy shocks, especially during the era of low inflation and deflation. Some studies report anomalous dynamics between variables that could be considered to be of importance, such as a "price puzzle" - a positive impact on prices of a contractionary monetary policy shock. Indeed, in the study by Sims (1992), the response of the price level to an interest rate shock was found to be positive. Similarly, most reported impulse responses in Jang and Ogaki (2003) show signs of a price puzzle. A policy proposal based on such results would be to increase interest rates in order to escape from deflation. Sims (1992) did not report confidence bands for the impulse responses, making it difficult to judge the statistical significance of the results. In Kim (1999), the responses of output and prices to a contractionary monetary shock have the expected signs; however, the error bands are sufficiently wide not to make either of the effects statistically different from zero. Moreover, the estimated benchmark systems by Miyao (2000, 2002) did not include a consumer price index or a GDP deflator, even if a consumer price index was included in alternative specifications as a robustness test. Finally, little evidence reaches far into the deflationary period. The samples of Bayoumi (2001), Miyao (2000, 2002) and Morsink and Bayoumi (2000) finish in 1998, Kim (1999) in 1996, Shioji (2000) in 1995, Kasa and Popper (1997) in 1994, Jang and Ogaki (2003) in 1993, and Sims (1992) in 1991. The longest samples are those of Kimura *et al.* (2002) ending in 2002 and Arai and Hoshi (2004) in 2003, the latter being comparable to our work. However, the study by Arai and Hoshi (2004) did not consider the potency of the interest rate channel; rather, the focus was on the existence of stable long-run relationships between narrow and broad definitions of the money stock and real output.

2.2.2 Monetary Policy at the Zero Lower Bound

We now turn from the existing evidence about monetary transmission in Japan to issues concerning the conduct of monetary policy during deflation or near the zero interest rate floor. Some of the aspects covered in this section pertain rather to the estimation of the structural IS equation in a later part of the paper. However, insofar as the setting of nominal interest rates is concerned, the VAR analysis can be considered relevant in this case as well. The importance of the money stock at the zero bound is also connected with the vector autoregressive approach, for we conduct estimations by using the money stock as an alternative indicator of the policy stance. This is pursued as an extension of the benchmark model where the short-term interest rate was considered to be the relevant policy variable.

Eggertsson and Woodford (2003), and Jung, Teranishi and Watanabe (2001) argued that a commitment to monetary easing in the form of keeping short-term interest rates low well into the future can make monetary easing effective even at the zero bound, as such a commitment can lower long-term interest rates. In particular, optimal policy was shown to be history-dependent, with monetary conditions remaining looser in the future even when the real deflationary disturbance has dissipated and the natural rate of interest has returned to a normal level. Expectations of a looser policy than one based on current economic conditions can then greatly eliminate the contractionary impact of the real disturbance. In the context of our empirical analysis based on the estimation of a structural IS equation, the existence of a stimulative effect of the real *ex ante* interest rate for output is interesting - the entire future paths of short-term real and nominal rates or very long term real rates were shown by Eggertsson and Woodford (2003) to matter for aggregate demand.

Such commitment solutions have advantages when compared to Taylor-type interest rate feedback rules in the context of the zero bound. Benhabib, Schmitt-Grohé and Uribe (2001) showed how Taylor rules could drive an economy into a liquidity trap through self-fulfilling paths of declining inflation rates and aggregate fluctuations. In the resulting state, the central bank is unable to pursue conventional interest rate policy to reverse a downward slide of prices. Rotemberg and Woodford (1999) and Williams (1999) considered inertial interest rate feedback rules instead:

$$i_t = (1-\theta)i^* + \theta i_{t-1} + \phi_\pi \pi_t + \phi_\chi \chi_t \qquad (2.1)$$

Here χ_t denotes the output gap, π_t is the inflation rate (deviation from target), i_t the nominal interest rate instrument of the central bank (and i^* the nominal rate when the real rate is at its long-run equilibrium level), with the coefficients on output gap and the inflation rate satisfying $\phi_\pi > 0$ and $\phi_\chi \geq 0$. According to such rules, interest rates are kept low even in the event of an increase in inflation or the output gap, simply because past interest rates have been low. This causes deflationary shocks to be followed by periods of higher than average inflation. Similarly, Reifschneider and Williams (1999) considered a rule of the following type:

$$i_t = \sum_{j=0}^{\infty} \theta^j [(\phi_\pi \pi_{t-j} + \phi_\chi \chi_{t-j}] \tag{2.2}$$

When the zero rate is not binding, the rules described in (2.1) and (2.2) are equivalent; when it does bind, interest rates remain low into the future in (2.2) because the central bank would have lowered interest rates in the past if the zero bound had not prevented it.

The commitment solution to the conduct of optimal monetary policy proposed by Eggertsson and Woodford (2003) was shown to be a version of a price level targeting rule. More generally, price level targeting as a monetary policy rule has drawn considerable interest since the increasing occurrence of deflationary episodes in various economies. This is true despite the fact that all inflation targeting central banks have currently opted for an inflation target instead of a (possibly upward trending) price level target, therefore treating inflationary shocks as bygones. In fact, only Sweden has ever officially pursued price level targeting (Berg and Jonung, 1999). An earlier view in the literature emphasized increased output and inflation variability as the result of a price level targeting approach (see Fischer, 1994). A theoretical study by Svensson (1999b) challenged this stance. The author found that with persistence in employment, price level targeting provides a 'free lunch' of identical output variability and reduced inflation variability, when compared to inflation targeting. Gaspar and Smets (2000) argued that price level targeting could actually reduce both inflation and output variability, conditional on forward-lookingness of economic agents and the credibility of the central bank. Moreover, due to the beneficial effect of expectations on real interest rates, the nominal rate has to adjust less, and the lower real *ex ante* interest rate can stimulate the economy after a period of disinflation. Indeed, as argued by Svensson (1999a), if a shock has pulled the price level below target, a credible commitment to a constant or an increasing price level target would automatically lower the real *ex ante* interest rate, even with no movement in the nominal rate. Price level targeting has also been an element of some other policy proposals, such as the "foolproof way" to escape from the liquidity trap presented by Svensson (2001, 2003). The strategy included the announcement of an upward-sloping targeted price path, and a currency devaluation with a subsequent peg to a crawling exchange rate target. When the price level target path has been reached, the currency peg will be abandoned in favour of flexible price level targeting with the former target path or flexible inflation targeting with the previous inflation target.[2]

[2]Interestingly, Coenen and Wieland (2004) found that switching to a price-level target

An emphasis on money supply instead of nominal interest rates may be interesting at the zero bound. Indeed, a monetary policy rule suggested by McCallum involves the management of the monetary base (see e.g. McCallum 1988, 2003). A focus on base money also has a role in the optimal interest rate policy of Eggertsson and Woodford (2003) as a signalling device of the commitment of the central bank. Even if the price level target was not actually met, the central bank would supply the amount of money that would be demanded if prices were at the targeted level. Furthermore, a focus on narrow money is consistent with the current operating target of the Bank of Japan. In March 2001, when the policy of quantitative easing was initiated, the current account reserves of commercial banks at the BOJ became the operating target for policy. The record-rate increases in the monetary base have yet to show in equal increases in broad monetary aggregates. This is evident in Figure 1 below, where we display the annual growth rates of monetary base and broad money M2+CDs.

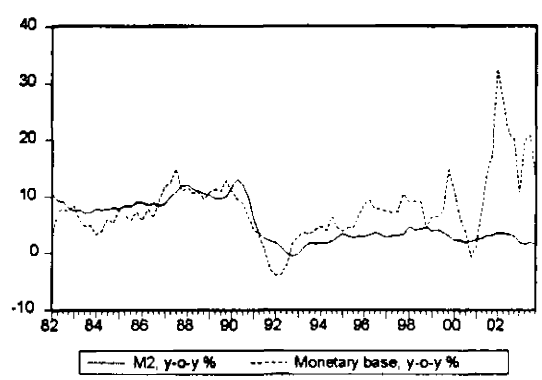


Figure 1. Annual growth rate of broad money M2+CDs and monetary base.

If broad money is a leading indicator for prices and output, and a stable relationship between broadly defined money stock and economic activity exists, monetary policy stance defined in terms of broad money growth may suggest a track record of contractionary, rather than expansionary, policy by the BOJ. This motivates the estimation of the VAR model with the broad money M2+CDs rather than the short-term interest rate as an indicator of the policy stance. Such considerations are further commented on in the methodology section of this paper.

alone is as effective in generating inflationary expectations as an exchange rate peg. In practice, however, an exchange rate peg may well be more credible, as this policy is easily verifiable by the public.

2.3 Methodology

An obvious problem with any empirical methodology for investigating the feasibility of a price level or an inflation targeting approach for Japan is that no such strategy was implemented in this country during the boom years of the 1980s, the "lost decade" of the 1990s or in the early years of the 2000s. Japan, however, is currently following a strategy that is very close to price level targeting, as a policy of quantitative easing was initiated in March 2001 in order to achieve a positive change in the price level. According to the most recent clarifications by the BOJ, the policy framework is to be in place until the CPI inflation rate stays at or above zero percent for a few months and there is no forecast by the BOJ board members of the economy falling back into deflation (Ito and Mishkin, 2004).[3] Even if we do not have data from a period of official price level targeting (or inflation, for that matter), this does not pose a significant limitation for the purpose of our study. Our main interest lies in the relationship between the interest rate, output and prices which of course matters for both inflation and price level targeting. The behaviour of economic agents with regard to the relationships between the investigated variables is probably only slightly affected by the change in the monetary policy strategy of the central bank, but it might be significantly influenced by a change from an inflationary to a deflationary environment. The Lucas-critique may then apply with force - a reduced form VAR essentially captures dynamic relationships between the variables, and is by construction a backward-looking representation of the data. Such considerations are tackled in the stability analysis of our VAR system. As a relatively stable model is established, we believe that the Lucas-critique does not pose a problem to the methodology chosen.

2.3.1 Vector Autoregressive Framework

The VAR-model can be written in its basic form similarly to Lütkepohl (2004):

$$x_t = A_1 x_{t-1} + ... + A_p x_{t-p} + CD_t + u_t \tag{2.3}$$

[3] The exact goals of the zero interest rate policy have slightly changed over time (Ito and Mishkin, 2004). When the emergency monetary policy was introduced in 1999, the intention was to continue until "deflationary concern is dispelled". Having briefly abandoned the zero rate policy in August 2000, only to re-introduce it in March 2001, the BOJ declared it would be in place until "the inflation rate becomes stably above zero". In March 2001, CPI excluding fresh food was identified as the relevant price index to focus on.

where p denotes the order of the VAR-model. K being the number of variables, $x_t = (x_t, ..., x_{Kt})'$ is a $(K \times 1)$ random vector, A_i are fixed $(K \times K)$ coefficient matrices and D_t contains all deterministic variables. Furthermore, C is a parameter matrix of a suitable dimension. The $u_t = (u_t, ..., u_{Kt})'$ is a K-dimensional white noise process with $E(u_t) = 0$.

As dynamic analysis within our benchmark VAR is based on a simple recursive ordering, our system could be estimated by OLS and the errors subsequently orthogonalized by a Cholesky decomposition. However, a structural form representation may be useful in the motivation of our identification scheme and the discussion of alternative (possibly nonrecursive) model structures. Ignoring the deterministic terms, a structural representation of (2.3) can be expressed as:

$$\mathrm{A}x_t = A_1^* x_{t-1} + ... + A_p^* x_{t-p} + \mathrm{B}\varepsilon_t \tag{2.4}$$

where $\varepsilon_t \sim (0, I_K)$. The matrix A allows for the modelling of the instantaneous relations, while B is a structural form parameter matrix. Again, the A_i^*'s $(i = 1, ..., p)$ are $(K \times K)$ coefficient matrices. The structural shocks ε_t are related to the model residuals by linear relations. Moreover, they are assumed to be mutually uncorrelated. Certain assumptions are necessary for identification, for it is not possible to directly observe the structural shocks. The reduced forms corresponding to the structural forms are obtained by multiplying (2.4) with A^{-1}, with $A_j = \mathrm{A}^{-1} A_j^* (j = 1, ..., p)$. We obtain the following relation between the reduced form disturbances and the structural form innovations:

$$u_t = \mathrm{A}^{-1}\mathrm{B}\varepsilon_t \tag{2.5}$$

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. the matrices A and B. The number of nonredundant elements of Σ_u, $K(K+1)/2$, is the maximum number of identifiable parameters in the matrices A and B. The overall number of elements in the structural form matrices A and B is $2K^2$. If either A or B is set as an identity matrix, we need to impose $2K^2 - K^2 - \frac{K(K+1)}{2} = K(K-1)/2$ additional restrictions to identify the full model. 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).

In order to examine the responses to Japanese monetary policy shocks, we estimate a simple benchmark trivariate VAR model ($K = 3$ in the framework above) including the (log) level of real GDP y_t, the (log) level of consumer prices (CPI) p_t, and the interest rate on the 3-month certificate of deposit i_t. The variables used in the estimation of the benchmark model are depicted in Figure 2 below. Regarding the choice of variables in the VAR, an alternative approach would have been to use the output gap, as is often done in monetary policy analysis and will indeed be the case in the estimation of the structural IS equation in a later part of this paper. However, using real GDP in levels ensures a common order of integration for all the endogenous variables in the VAR estimation. It is appropriate to consider the short-term nominal rate in order to capture the monetary policy shocks, as there is considerable consensus in the literature that it is a good indicator of the policy stance. Moreover, the short-term operating target in Japan has historically been the short-term interest rate (see e.g. Okina, 1993; Ueda, 1993). Admittedly, the 3-month rate is not the policy rate of the Bank of Japan. But importantly, using the rate on the 3-month certificate of deposit allows us to include variations in the nominal interest rate in our vector autoregressions even after the central bank's main policy rate hit zero, with our rate likely to better reflect the actual cost of lending to households and firms than the measure used in the conduct of the zero interest rate policy. Indeed, the call rate fell to 0.03% already in the second quarter of 1999 whereas the rate on the 3-month certificate of deposit only hit a correspondingly low level of 0.04% in early 2001.[4]

[4]Similarly, Bayoumi (2001) used the 3-month Gensaki rate in his study (even if the variable used in the estimation was the *real* interest rate constructed by using the nominal 3-month rate). In contrast, Miyao (2002) considered the call rate that was obtained by combining the uncollaterized and collaterized rates, with an adjustment to the latter. Morsink and Bayoumi (2000) used the uncollaterized overnight call rate.

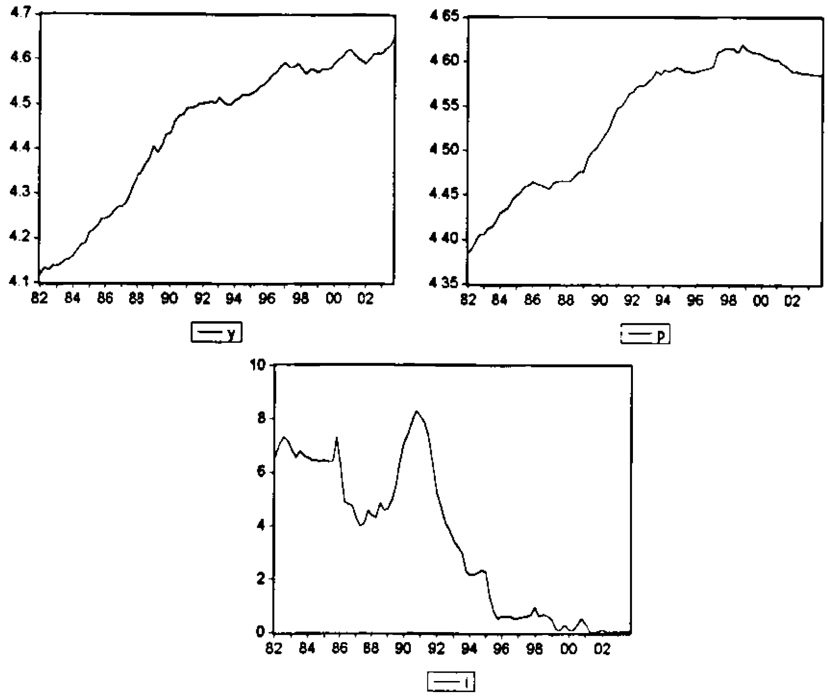


Figure 2. Series used in benchmark VAR estimation. Real GDP (y), consumer price index (p) and interest rate on 3-month certificate of deposit (i). Output and prices in logarithms.

As a benchmark model, we consider a simple recursive identification scheme, whereby we can write $A = I_K$ and $u_t = B\varepsilon_t$. When we restrict the B matrix to be lower triangular, the first component of ε_t, ε_{1t}, can have an instantaneous impact on all three equations, whereas the second component ε_{2t} can have an effect on all the other equations except the first, and so on. With the ordering of the variables written as y_t, p_t and i_t, the interest rate (real output) can then be considered the most (least) responsive to changes in economic conditions. Importantly, neither the price level nor output react contemporaneously to a monetary shock (defined as a shock to the nominal interest rate) - a canonical assumption in monetary policy analysis. In addition to its intuitive appeal and simplicity, such an identification scheme can be theoretically motivated. The model by Rotemberg and Woodford (1999) implies that the interest rate decision at period t cannot affect the determination of period t output or inflation (price level in our case), as interest-sensitive private expenditure is to some extent predetermined. As a consequence, a positive shock to the nominal interest rate truly represents a monetary contraction, as inflation (the price level) is unchanged during the period of the shock. Of course, the nominal interest rate may quickly return to its pre-shock level. But as emphasized by Woodford (2003), if inflation is lowered for several quarters, a persistently higher real interest rate is the source of the monetary contraction - it is the real interest rate rather than the nominal one that matters for aggregate demand. In their analysis of policy rules for inflation targeting, Rudebusch and Svensson (1999) discussed how their structural model with an AS equation and an IS curve could be interpreted as a VAR model with a Cholesky decomposition. The causal order of the variables was output, inflation and the interest rate, identically to our system. The AS equation was specified to be a Phillips-curve with autoregressive (adaptive) expectations - therefore a backward-looking model was established to examine various inflation targeting rules.

Our identification scheme assumes that the monetary authority observes y_t and p_t when choosing i_t, which may be debatable. In Japan, information on the consumer price index is rather timely, available 28 days after the reference month, while information on GDP is only available in the middle of the second month after each quarter. Yet more timely data on economic activity is available in the form of monthly industrial production data within one month of the end of the reference month. It is generally assumed that the monetary authority considers all the information at hand when taking a decision regarding the policy stance. Moreover, by imposing output to affect the price level and the interest rate setting contemporaneously, we assume that economic activity bears information about prices that the central bank takes into account in its policy setting. Nevertheless, we discuss possible alternative identification schemes in the context of model estimation that deviate from the benchmark recursive ordering.

Note that if the third row of the B matrix is considered a monetary policy reaction function, the inclusion of the price level instead of the inflation rate would make it resemble a Wicksellian interest rate rule defined by Woodford (2003) as:

$$i_t = \phi(P_t/P_t^*; v_t) \tag{2.6}$$

where P_t^* defines a target path for the price level and v_t is a possible random disturbance to the policy rule or to its implementation. The function ϕ indicates the rule used by the monetary authority in setting its operating target. As the price-level target P_t^* can vary over time, we can consider the rule either as one stabilizing the price level around a growing trend path, or a constant price level target. Moreover, the central bank is allowed to respond to output variations as well as to the price level path - this is implied by the random disturbance term that could capture such exogenous influences.

In addition to the benchmark VAR model described above, we consider a trivariate model replacing the short-term interest rate by the broad money stock M2+CDs as the monetary policy indicator, and examine the predictive content of various components (both narrow and broad) of the money stock with the use of Granger-causality tests. While theoretical feedback rules could be defined in terms of the monetary base where money supply is a specified function of the price level, an increase in broad money supply may be difficult to bring about in the case of instabilities in the money multiplier. Acknowledging such problems, we nevertheless replace the interest rate by broad money in the B matrix. This effectively leads to considering a feedback rule in terms of broad money, where shocks to broad money have no contemporaneous impact on output and prices, whereas money supply is contemporaneously affected by both output and price shocks. West (1993) uses a similar rule for the Japanese M2, interpreting his results as consistent with a money supply rule for a central bank that aims at output and price stabilization.[5]

Our interest in a broad monetary aggregate as an indicator of policy stance is driven by empirical rather than theoretical motivations, however. Firstly, as reported by Miyao (2002), a strand in the Japanese monetary policy literature views the money stock as an important indicator of policy. While M2+CDs never obtained the status of an intermediate target, it was declared to be an important leading indicator for prices as early as 1975, and projections for the growth rate of M2 (later M2+CDs) over the coming quarter ("quarterly foresights") were produced from July 1978 onwards. Secondly, for Japan a decline in short-term policy interest rates - the usual source of monetary stimulus - has clearly been unavailable since 1999. Thirdly, as mentioned in the introduction, some theoretical considerations focus on the management of the monetary base at the zero bound, consistently with the current operating target of the Bank of Japan. Finally, comparisons between various components of money in the analysis allow us to investigate the implications of monetary expansion, both narrowly and broadly defined.

[5]Interestingly, Morsink and Bayoumi (2000) found that the Japanese bubble in the late 1980s was driven by broad money and loan shocks rather than by interest rate shocks. This could be seen to justify the interest in broad money as an indicator of policy stance. Indeed, Lipworth and Meredith (1998) suggested that interest rates provided a misleading picture of the policy stance in the late 1980s and early 1990s. In a similar vein, Ueda (1993) found that monetary aggregates are leading indicators for the real economy, while such a role is not clear for interest rates in Japan.

It would seem a bold claim to suggest that no other variables would be important in the Japanese context, given the problems of non-performing loans, interventions in the foreign exchange markets, and the controversy surrounding arguments about using expansionary fiscal policy even in an environment of spiralling public debt.[6] However, the trivariate model captures the main variables of interest in the monetary transmission mechanism. Moreover, small-dimensional systems are often found to perform well in time-series frameworks, especially given the low power of cointegration tests and the large number of (often statistically insignificant) parameters in high-dimensional vector autoregressive or vector error correction models. Finally, it seems reasonable to suggest that in our relatively short sample only a system with a small enough number of dimensions would yield precise and reliable estimates.

2.3.2 Data Sources

We use quarterly, seasonally adjusted data from the OECD, IFS and the Bank of Japan Databases for the period 1982Q1-2003Q4 in the VAR estimation. The precise data sources are listed in Appendix A. All variables except the interest rate were transformed into logarithms. In the VAR estimations, we predominantly used the software JMulTi (2004), version 3.11. Exceptions are mentioned in the text. EViews was used in the estimation of the structural IS equation in the later part of this paper. Our choice of sample period allows for the examination of the monetary transmission mechanism during a period of positive inflation rates and the booming economy of the late 1980s and the early 1990s, and the subsequent disinflation that turned to deflation in 1995.[7] The Japanese economy was in a slump for most of the 1990s, with an average growth rate during 1993-2003 of slightly above 1 percent. A meagre recovery took place from 1995 to 1997 prior to the pre-announced consumption tax increase from 3% to 5% of April 1997. It may be important to explicitly take this particular tax hike into account, as it was introduced in the middle of the low inflation period and induced some output fluctuation in the economy.

[6] For a discussion of the possible Japanese credit crunch and banking sector problems, see e.g. Bayoumi, 2001; Kato *et al.*, 1999; Woo, 1999; Kashyap, 2002; for proposals to use the exchange rate channel to induce a positive inflation rate in Japan, see e.g. Coenen and Wieland, 2003; Fujii, 2004; Svensson, 2001; for arguments about the conduct of fiscal policy in the Japanese case, see e.g. Posen, 1998.

[7] Moreover, there are visible outliers in the series of consumer prices and the nominal interest rate still in 1980, motivating a starting period of 1982 in the VAR estimation.

Accordingly, a shift dummy was included in the estimation, taking the value of zero prior to 1997Q2 and 1 from 1997Q2 onwards.

2.3.3 Unit Root and Cointegration Tests

In order to determine the order of integration of the series, we performed the augmented Dickey-Fuller (ADF) test for all the series. As our approach, we used the so-called Pantula principle (see Pantula, 1989), whereby the series is initially differenced a sufficient number of times in order to make it stationary. If the unit root is rejected for this series, the test is performed on the series differenced one time less than previously. This procedure is repeated until the null hypothesis of unit root cannot be rejected.

We commenced the ADF-testing with all the series in second differences, assuming a maximum order of integration of 2 for the series. The results from the unit root tests are listed in Appendix B. Using a graphical inference in order to determine the appropriate deterministic terms to include in the testing procedure, we found strong evidence that real output would be an I(1) variable. We were able to reject the null hypothesis of a unit root for the price level series in first differences, even if only at 10% level. The estimated models performed satisfactorily when treating the price level as integrated of order one, and we therefore continued with the assumption that the price level in first differences is stationary.[8] The nominal interest rate was found to be an I(1) variable (even with the zero bound!) and was treated as such in the estimation, again due to the satisfactory performance of our estimated model with such an assumption.

Given our knowledge that all variables in the benchmark VAR could be characterized as I(1), we moved on to examine the cointegration properties of the endogenous variables. The results from the cointegration tests are displayed in Appendix B. The Saikkonen-Lütkepohl test (Saikkonen and Lütkepohl, 2000a, b) was used for the purpose of examining the existence of common stochastic trends. The test first estimates the deterministic term by a GLS procedure, then subtracts it from the observations and applies a Johansen type of test to the adjusted series. The latter test is a likelihood ratio

[8]Surprisingly, when a unit root test with a structural break (see Lanne *et al.*, 2002) was used in order to account for the consumption tax hike of 1997Q2, a unit root could not be rejected for the price series in first differences even at 10% level. Setting the break date to any other quarter in 1997 would bring about a rejection of a unit root, however.

test based on a reduced rank regression of a vector error correction specification (Johansen, 1995). Little robust evidence was found in the cointegration testing, as the results were rather dependent on the lag length, the deterministic terms used, and the starting date of the sample. When a constant and the shift dummy variable for 1997Q2 were included as the deterministic terms, we were actually able to reject the cointegration rank of 2, suggesting that all the variables were stationary, I(0). Such a conclusion could be rejected on the basis of the unit root tests, however. An identical finding arose in the Johansen trace test, when Hannan-Quinn or Schwarz criteria were used to determine the lag length. When a linear trend was additionally included as a deterministic term, as may be preferable due to the trending behaviour of the series, we in contrast obtained evidence in favour of one cointegrating relationship in the Saikkonen-Lütkepohl test. The assumption of an orthogonal trend even led to the rejection of a cointegration rank of one. However, the finding of a cointegration rank of one obtained with a fully general linear trend breaks down when the sample size is increased by just a few observations: when the start date for the sample was moved from 1982Q1 to any date between 1980Q1 and 1981Q1, we were no longer able to reject the null hypothesis of no cointegration in our system.[9]

Given the low power of cointegration tests in our relatively small sample, and the fact that our interest is not centred on possible cointegration relationships, we opted for estimating a VAR model where all the endogenous variables are represented in levels. Such an approach avoids a misspecification problem in the case where cointegrating relations between our variables actually existed. Furthermore, as we are estimating a system with a lag length of 2 or higher, the usual tests and t-values have their asymptotic properties, allowing for statistical inference (Dolado and Lütkepohl, 1996).

2.4 Empirical Evidence

In this section, we present the estimation results. We commence with the evidence yielded by vector autoregressions, using both the nominal interest rate and broad money as indicators of the monetary policy stance. We then discuss the policy implications that could be derived from the estimations of

[9]This finding holds when the Akaike criteria are used to determine the lag length for 1980Q1-1981Q1; similarly with the Hannan-Quinn and Schwarz criteria for 1980Q2-1980Q4.

such models. Finally, results from the structural IS equations are presented and commented upon.

2.4.1 Estimation of the Benchmark VAR Model

The reduced form VAR model was estimated for the period 1982Q1-2003Q1, with the vector of dependent variables for the benchmark model written as $x_t = (y_t, p_t, i_t)'$. The lag order was chosen to be 3 on the basis of misspecification and stability tests, coupled with the consideration of having an adequate number of lags to properly examine the monetary transmission mechanism. This choice of lag length also determined that the observations 1982Q1-1982Q3 were used as presample values, leaving us with a total of T=85 observations. A constant, trend, and a shift dummy taking a value of one from 1997Q2 onwards were included as deterministic terms, and the system was estimated by OLS.

Our reduced form model was submitted to various misspecification tests, as displayed in Appendix C. We performed the Portmanteau and LM-tests for residual autocorrelation, multivariate Jarque-Bera and single equation ARCH-LM tests to detect nonnormality and ARCH-effects in the residuals, respectively. All in all, the residual testing did not give rise to a rejection of the model, even if there is evidence of kurtosis in the VAR. However, since the asymptotic properties of VAR-estimators are not dependent on the normality assumption, this point may be of minor importance for our purposes.

Economic inference is difficult and possibly misleading from the large number of autoregressive coefficients of a reduced form model, for the system is essentially an atheoretical representation of the dynamics between the variables. Some points regarding the reduced form representation are worth mentioning, however. The coefficients on the first and second lags of the interest rate variable in the third (interest rate) equation of the system - both statistically significant - add up to 0.78. Such a finding could be regarded as suggesting considerable interest rate smoothing by the Bank of Japan. Rudebusch and Svensson (1999) claimed that gradualism in interest rate setting is characteristic of all inflation-targeting central banks. One could even argue that the Bank of Japan was indeed an implicit inflation targeter with a zero inflation target - in October 2000 the BOJ issued a report "On Price Stability" where price stability was defined as a state that is neither inflation nor deflation (Bank of Japan, 2000). Perceived in such a way, persistent deflation would then have merely been the outcome of some (rather serious) policy mistakes. Mishkin (1999) suggested that the inclusion of interest rate variability in the central bank loss function may reflect concern about financial stability. In the Japanese context, it is likely that the persistence of interest rates near or at the zero floor may have made debt-servicing feasible for many firms with financial problems.

A scoring algorithm to obtain the estimates for the structural form B matrix yields the following outcome, with the standard errors in parentheses:

$$\begin{bmatrix} \begin{matrix} 0.0083 \\ (0.0006) \end{matrix} & 0 & 0 \\ \begin{matrix} -0.0006 \\ (0.0004) \end{matrix} & \begin{matrix} 0.0034 \\ (0.0003) \end{matrix} & 0 \\ \begin{matrix} 0.0881 \\ (0.0350) \end{matrix} & \begin{matrix} 0.0655 \\ (0.0340) \end{matrix} & \begin{matrix} 0.3096 \\ (0.0237) \end{matrix} \end{bmatrix} \tag{2.7}$$

While the point estimates of the contemporaneous impulses may not be worthy of note for the present purposes, their signs could be of interest. If the third row of the above matrix is considered a type of monetary policy reaction function (or an instrument rule with two target variables), the coefficients b_{31} and b_{32} on real output and prices, respectively, obtain the expected signs. More information can be derived from an impulse response analysis of the structural form model. In order to construct the confidence intervals to illustrate parameter uncertainty, we used the bootstrapped Hall 95% percentile confidence interval, with the number of bootstrapping replications set at 1,000 (see Benkwitz *et al.*, 2001). In addition to the advantage of enhanced small sample properties of bootstrap confidence intervals in comparison with other asymptotic methodologies, this particular approach benefits from a built-in bias adjustment. All the impulse responses of the system are displayed in Figure 3 below.

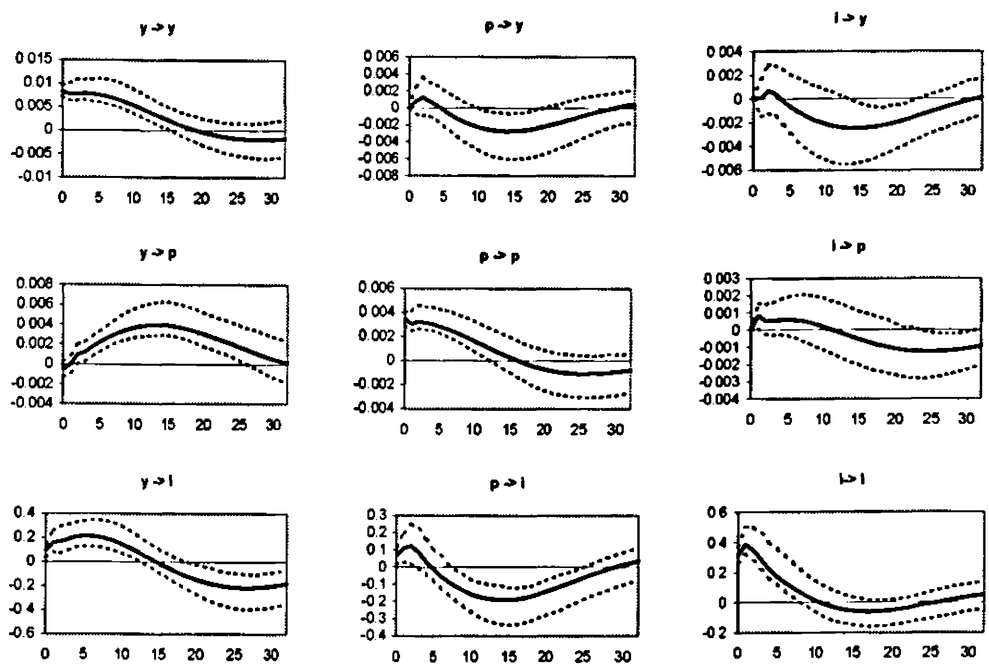


Figure 3. Impulse responses of real GDP (y), price level (p), and interest rate (i) to output shocks (first column), price shocks (second column) and monetary policy shocks (third column).

According to our estimated impulse responses, a positive interest rate shock leads to a significant fall in real output, suggesting the existence of a potent interest rate channel in the Japanese economy. Such a result is consistent with standard Keynesian models with wage-price inertia. The multiplier, calculated as the maximum impact on real output of the peak in the interest rate shock, amounts to -0.6.[10] In line with conventional wisdom about the effects of monetary policy shocks, the impact on output is transitory. Furthermore, an increase in the nominal interest rate also leads to a fall in the price level in the long run, even if (probably due to inflation persistence) the price level initially rises after the interest rate shock in our system. The stimulative impact on real output also holds for the real interest rate; an investigation of the accumulated impulse responses, together with

[10]This is obtained as a fall in output by 0.25 percent as a result of a 39 basis point increase in the nominal interest rate. Interestingly, the multiplier is identical to the one found by Morsink and Bayoumi (2000) for 1980-1998. The authors used real private demand (real GDP minus total government spending) as a ratio to potential output, while the interest rate was the uncollaterized overnight call rate. The peak impact on output was obtained after 8-10 quarters, slightly faster than in our study.

the impulse response coefficients, reveals that we are indeed dealing with a monetary contraction defined in terms of the real interest rate. The responses of both prices and output to a monetary policy shock are sluggish and, as is generally assumed, the impact on prices is slower than that on output.

Even if the above dynamics (at least their signs) seem to be consistent with common perceptions about the effects of monetary policy shocks, rather strong words of caution are in order against hasty conclusions in favour of price level or inflation targeting strategies. Firstly and importantly, note the very slow impact of monetary policy shocks on the price level. The price level only starts to fall below the initial level after 3 years have passed from the shock, while it reaches its lowest point only after 6 years! In Shioji (2000), prices were found to fall somewhat faster, after 2 years approximately, but a statistically significant effect took considerably longer to come about (in fact, using two standard error bands, the effect on prices was never statistically significant). Similarly, we find that output is rather slow to react to a monetary policy shock, with the peak experienced only after 3 years have passed from the shock. This is consistent with the finding by Miayo (2002), whose results seem to show a peak impact at around 30 months. In Bayoumi (2001) where the real interest rate was used, the peak impact on output occurred significantly faster, after 6 quarters. Overall, our results confirm the empirical observation that deflation seems deeply entrenched in the Japanese economy; even the recovery of late 2003 with its high GDP growth rates seemed unable to do much to induce a positive change in the price level. Nevertheless, the stimulative (though slow) effect of a monetary policy shock is a significant finding, given the number of years in our sample when deflation actually prevailed in the Japanese economy and the minimal movements of the nominal interest rate close to the zero lower bound during that period.

We proceeded with tests to examine the stability of our estimated model. These are of importance in our case, as instability in the system could provide information about possible structural breaks in the monetary transmission mechanism during the transition of the Japanese economy from an inflationary to a deflationary era. Results from the model residual-based CUSUM of squares test were taken at suggestive value only, as the validity of CUSUM type tests in systems with integrated variables is not clear (see Lütkepohl, 2004). The test statistic never crossed the critical lines in any of the model equations at a 1% significance level; furthermore, only the output equation showed weak signs of instability at a 5% level. These results are available upon request.

Further stability testing was performed by the Chow forecast and sample split tests, where the null hypothesis of time invariant parameters was tested against the possibility of a change in the parameter values at a certain date. These results are displayed in Appendix D. As suggested by Candelon and Lütkepohl (2001), the approximate χ^2 and F-distributions of the Chow test statistics can be rather poor approximations and may lead to very high rejection rates, especially in small samples.[11] This was also the finding in our stability tests. We therefore used bootstrap p-values, based on 2,000 replications, and tested every third observation within a sample of 1993Q1-2003Q3 for model stability. Had a break in the transmission mechanism actually taken place, such an event would have probably occurred during this time period. Indeed, Miyao (2000) suggested a break date in 1995; similarly, examining the relationship between broad money M2+CDs and output, Arai and Hoshi (2004) suggested a break date of late 1995. The latter authors also found a break date for a cointegration relationship between M2+CDs and monetary base in 1998. Additionally, it is for this period that the CUSUM of squares test showed weak instability in the output equation. Using the Chow forecast test, the stability of our estimated system could never be rejected at conventional significance levels. With the sample split test, a low p-value was detected for a single observation only.

It may be of additional interest to test every observation for a structural break during 1995-1997; at this time, the short-term interest rates were lowered to record-low levels (the official discount rate was lowered to 0.5% in September 1995), and the financial sector problems culminated in the failing of two big financial institutions, the Hokkaido Takushoku Bank (the tenth largest city bank at that time) and Yamaichi Securities (the fourth largest stockbroker) in November 1997.[12] With this testing sample, model stability was never rejected in either of the two Chow tests at conventional levels of significance. Additionally, we considered the recursive fluctuation tests proposed by Ploberger *et al.* (1989) for each of the three equations.[13] The tests were bootstrapped with 4,999 replications, and p-values from the resulting empirical distributions were used accordingly. For none of the three equations do we find evidence of instability, as illustrated in Appendix D. Finally, we es-

[11] The authors simulated the stochastic part of a monetary model for Germany in their Monte Carlo tests. It was found that the original Chow tests had unacceptable rejection frequencies even in sample sizes of $T = 300$. The performance of the bootstrap versions of the tests was found to be much better even for a smaller sample size of $T = 76$.

[12] In August 1995, both the largest credit union (Kizu Credit Union) and the largest secondary regional bank (Hyogo Bank) had already failed, further suggesting a possible break date at some time during 1995-1997 (Miyao, 2005).

[13] This test was performed using the Structural VAR program by Warne (2005).

timated the model until 1995Q4 only and compared the previously obtained peak responses of output and prices to monetary policy shocks with those resulting from the shorter estimation sample. All of the previously obtained point impacts fall within the new 95% confidence intervals. We therefore concluded that the majority of our evidence points to a stable model.

VAR estimations in the Japanese context have been predominantly performed in a linear framework, as has also been the case above. To our knowledge, only Kimura *et al.* (2002) have used a Bayesian VAR with time-varying coefficients. Nevertheless, there could be some justifications for doubting the validity of a linear approximation to studying monetary policy near the zero interest rate floor. This primarily concerns the nominal interest rate that could be thought to show nonlinear behaviour when approaching the lower bound. To investigate the adequacy of our linear model, we performed linearity tests as proposed by Teräsvirta (1994), where the linearity hypothesis was tested against a smooth transition regression model. Here we fully acknowledge the caveat that with integrated variables, the validity of the tests may be problematic. The tests were performed on each of the three equations of the VAR model, where the dependent variable was characterized as endogenous and the remaining two were set as exogenous. Unrestricted tests, however, suffered from the near singularity of the moment matrix. We therefore set the deterministic terms to only appear linearly; similarly, whenever output and prices were used as exogenous variables these were also restricted to appear linearly. The latter restriction can be justified on the basis of the assumption that it is indeed the interest rate that displays the nonlinear behaviour. Accordingly, the nominal interest rate was used as the transition variable from one regime to another. With such restrictions, the entire testing sequence described in Teräsvirta (1994) was plausible. The results from the linearity tests are illustrated in Appendix E. For none of the three equations could we reject the null hypothesis of linearity, for any of the three lags of our transition variable. It therefore appears that our linear model is adequate for the purpose of our study. Such an outcome is consistent with the finding of relative stability and a rather satisfactory outcome from the misspecification tests, even if some kurtosis was detected in normality testing.

We considered various alternative structures for the benchmark model, given the easy adaptability of a structural VAR with contemporaneous restrictions to slight modifications in the identification scheme. It was mentioned in the methodology section that the timing assumption with regard to the information available for the monetary authority at the time of the interest rate decision in period t is arguable. Specifically, information on real GDP is not available for the Japanese monetary authority at the current quarter (even if industrial production data is still at hand). Such considerations could motivate an over-identifying restriction of setting the contemporaneous impact of real output on the interest rate to zero (coefficient b_{31} in the B matrix below), while the matrix A is set as an identity matrix as before. This results in the following matrix for B:

$$\begin{bmatrix} b_{11} & 0 & 0 \\ b_{21} & b_{22} & 0 \\ 0 & b_{32} & b_{33} \end{bmatrix} \tag{2.8}$$

Our restriction, however, was rejected by the data at a 5% level. A formal likelihood ratio test yields a statistic of $\chi^2 = 6.195$, with a p-value of 0.012. We then considered the contemporaneous identification scheme proposed by Sims (1992) whereby monetary policy shocks affect all other variables contemporaneously, while shocks to output do not contemporaneously affect any other variables. Such an ordering violates the consideration of interest setting as a feedback rule for (contemporaneous) prices and output. Moreover, on the basis of empirical regularities and theoretical considerations presented in the methodology section, it is difficult to argue that the impact on prices and output of monetary policy shocks could come about within the same quarter. Such a model structure could nevertheless be considered a robustness test. In the context of our variables, this gives the following matrix for B:

$$\begin{bmatrix} b_{11} & b_{12} & b_{13} \\ 0 & b_{22} & b_{23} \\ 0 & 0 & b_{33} \end{bmatrix} \tag{2.9}$$

The above ordering produced materially similar dynamics to our benchmark system (excluding the contemporaneous impacts of course), even if the significance of the interest rate shocks on prices declined. As an additional robustness test, we shortened the estimation period by one year, both from the beginning and the end of the sample, resulting in the periods 1983Q1-2003Q4 and 1982Q1-2002Q4. Both samples create results that are qualitatively very close to the benchmark estimation; moreover, the signs of the coefficients on the monetary policy rule of the B matrix remain unchanged. Finally, we replaced the consumer price index by the GDP deflator. With this price indicator the significance of monetary policy shocks in the system was somewhat weakened. These results are available from the author upon request.

2.4.2 Models with Money

In addition to the benchmark model described in the previous section, we considered a trivariate system where the nominal interest rate was replaced by a broad monetary aggregate, M2+CDs (in real terms, deflated by the consumer price index). This measure of broad money consists of notes and coin in circulation, sight and time deposits, and certificates of deposit. Regarding the time series properties of the broad monetary aggregate, unit root tests with a structural break in 1990Q2 by Lanne *et al.* (2002) rejected the unit root for the broad money series in first differences. Such a break date is consistent with the bursting of the Japanese financial bubble that induced slower growth rates for broad money. Results from the Saikkonen-Lütkepohl test did not provide evidence in favour of rejecting a cointegrating rank of zero in a system including the three endogenous variables, a constant term and dummy variables. Including a trend term would bring about a suggestion of a cointegrating rank of one; however, this result was not robust to the omission of dummy variables. Moreover, the assumption of an orthogonal trend leads to the nonrejection of a cointegrating rank of zero. These results are available upon request.

All variables were included in the VAR in first differences (in stationary form) with 3 lags. As the possibility of cointegration between the variables could not be ruled out, a misspecification issue arises that could be avoided by estimating the system in levels. However, misspecification tests for residual autocorrelation afforded somewhat unsatisfactory results in a levels form VAR including broad money. The deterministic terms included a constant, trend, an impulse dummy for 1997Q2 (corresponding to the shift in the benchmark levels VAR), and an additional shift dummy for the period 1990Q2-1992Q2 to account for the decreasing growth rates in broad money at the time of the bursting of the asset price bubble. The estimation of a stationary system allows us to conduct a model reduction procedure, whereby at each step the coefficient with the lowest t-value in the entire system was checked and possibly eliminated. Only coefficients with a higher t-value than the specified threshold of 1.67 were then eventually maintained. Such a procedure can be considered useful in VAR models that typically feature a large number of statistically insignificant coefficients.

Our model where monetary shocks are now identified as shocks to M2+CDs, passes all specification tests, as indicated in Appendix C. Moreover, as CUSUM type tests are valid for an investigation of model stability due to stationarity of the underlying variables, they are of interest for our purposes. The test statistic never crosses the critical lines in the CUSUM of squares test at a 5% significance level, suggesting a stable model. Conducting the Chow sample split and forecast tests in an identical fashion to the benchmark model with interest rates, parameter constancy cannot be rejected for any tested break date. The stability test results are illustrated in Appendix D.

The model where restrictions on the B matrix are specified identically to the benchmark model above was estimated by a scoring algorithm. An investigation of system dynamics in the context of impulse response analysis reveals that an increase in real M2+CDs leads to a significant increase in both the price level and output, as illustrated in Figure 4 below. The response of the first differences in prices and output has been accumulated to illustrate the impact on the levels of the corresponding variables in the graphs. Our finding is of interest, given the persistence of the deflationary environment in Japan and the apparent ineffectiveness of monetary policy in pulling the economy out from a deflationary trap. The results are consistent with the outcome of the benchmark model, where shocks to an alternative measure of the monetary policy stance were found to significantly affect both output and the price level. Morsink and Bayoumi (2000) also found shocks to broad money to lead to increases in both private demand and the price level. Similarly, while Arai and Hoshi (2004) focused on the long-run relations between M2+CDs, real GDP and monetary base, they found that a positive long-run relationship between M2+CDs and real GDP, and between M2+CDs and monetary base, still existed in their estimation sample that ended in the last quarter of 2003. Measured in this sense, monetary policy had not lost its effectiveness.

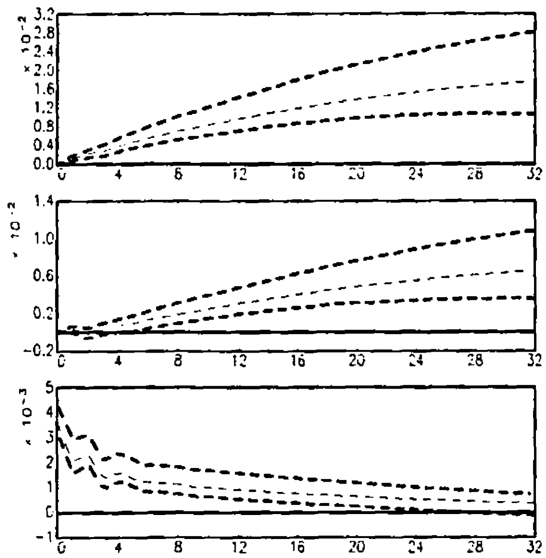


Figure 4. Impulse responses of real GDP (top), price level (centre), and M2+CDs (below) to broad money shocks. Response of M2+CDs in first differences.

Would an expansion in narrow as well as in broad money be of equal help for the Japanese authorities in creating positive inflation? To answer this question, Granger causality tests were conducted between the variables in our model, including real output, the consumer price level, and one component of the money stock. Different components of money were tested in order to determine possible differences in their causality properties with respect to output and prices. According to the definition of Granger causality, a variable x_t is causal for a time series variable y_t if the former helps to improve the forecasts of the latter. Then, causality is determined by a variable's forecasting ability - this is admittedly rather different from economic causality. However, considering money an indicator of the monetary policy stance, such causality could yield useful information about the future price path of a deflationary economy. We divide our trivariate system with its vector of three variables x_t into two subvectors, x_{1t} and x_{2t}, where x_{1t} includes the monetary aggregate in question and x_{2t} contains real output and the consumer price index. In this setup, in a partitioned VAR of order k, x_{1t} is Granger non-causal for x_{2t}, if and only if $\alpha_{21,i} = 0$, for $i = 1, ..., k$. The null hypothesis of noncausality is rejected if at least one of the $\alpha_{21,i}$ is nonzero. The usual definition of Granger-causality can therefore be applied to the two subvectors of the vector x_t including all the variables of the system, while it is clear that we cannot imply that a member of one subsystem (say, money), is causal for *each* of the variables in the other subsystem (prices or output). Standard χ^2- and F-tests based on Wald principle could be considered when all the variables are stationary. However, we include all the variables in levels. Relevantly for such a case with integrated variables, Dolado and Lütkepohl (1996) suggested that when some or all of the variables are integrated of order $d > 0$, the Wald-test can be modified in order for it to have a χ^2 -distribution. Accordingly, we conducted a lag inclusion test to examine whether the coefficient matrix $A_{p+1} = 0$. The Granger-causality test was subsequently performed on the A_i, $i = 1, ..., p$, only, ignoring the redundant lag. The results from the Granger-causality tests are displayed in Table 1 below.

	H$_0$: z does not Granger-cause x	
Series	Test statistic (χ^2)	p-value
z=broadly defined liquidity, x=gdp, cpi	16.11	0.04**
z=M2+CDs, x=gdp, cpi	17.88	0.00***
z=M1, x=gdp, cpi	1.54	0.99
z=monetary base, x=gdp, cpi	12.62	0.13
z=notes in circulation, x=gdp, cpi	4.50	0.34

Table 1. Granger-causality tests between components of the money stock, consumer price index and real output.

Note: * indicates significance at 10% level, ** at 5% and *** at 1% level.
Shift dummy for 1997Q2 included in all systems.
Additional shifts 1990Q2-1992Q2 for systems with M2+CDs and broadly defined liquidity.
Lag length without the redundant lag: 4 for monetary base and M1;
4 for broadly defined liquidity; 2 for M2+CDs and notes in circulation.

The results are remarkably consistent within the groups of narrow (represented by monetary base, M1 and notes in circulation) and broad (M2+CDs and broadly defined liquidity) components of the money stock. Firstly, all indicators of broad money were found to be causal for real output and prices at a 5% significance level. Secondly, for none of the components of narrow money could we reject non-causality at conventional significance levels. As causality tests are generally based on fairly large models, they may have weak power in rejecting the null of noncausality. In that regard, our results of the broad monetary aggregates being causal for the price level and real output appear rather strong. The findings from the Granger-causality analysis can then be seen as providing further support for our VAR results indicating beneficial effects of broad money expansion. Our results are similar to West

(1993) who found M2 to Granger cause both output and prices during 1973-1990; the author interpreted the results as suggesting that monetary policy could then be used to influence the path of the latter two variables.[14] Interestingly, high-powered money was reported not to have the same property, in accordance with our results. As pointed out before, however, the increases in broad money are difficult to bring about with recent instabilities in the Japanese money multiplier. Finally, for M1 and M2+CDs, we witness significant causality from output and prices to money (results not reported), possibly due to changes in the transactions demand for money as a result of changes in real output or the price level.

It would be easy to argue that our results of narrow money not being Granger-causal for prices and output would be in line with recent developments in the Japanese economy. Specifically, quantitative easing with its implied increases in monetary base (whose annual growth rate exceeded 30% in the first quarter of 2002!) seems to have had little impact on prices. However, it is not clear what the outcome in terms of output and prices would have been if a policy of quantitative easing had never been implemented. By providing ample liquidity the Bank of Japan can help the recovery by stabilizing the financial markets; such a course of action is likely to have an indirect positive effect on the price level.

2.4.3 Policy Implications

Our previous results from the benchmark model established the effectiveness of the interest rate channel in the Japanese economy, even during an era when inflation rates have moved from positive to negative territory. Our findings may be considered to justify a commitment policy to keeping nominal interest rates low well into the future to stimulate aggregate demand, as proposed by Eggertsson and Woodford (2003), and Jung, Teranishi and Watanabe (2001). Similarly, as the results were obtained in a system such as our VAR where no forward-looking variables are included, they could be regarded as advocating a policy that sets an inflation target for the central bank - inflation targeting can effectively be pursued in a backward-looking setting also. However, the transmission of monetary policy shocks to the real economy appears to be

[14] Towe (1998) found in bivariate Granger-causality testing that monetary aggregates were able to predict real activity well in the 1990s, while the forecasting ability of nominal interest rates fell during this period. The author suggested that attempts to meet the BIS capital adequacy requirements and balance sheet problems affected credit conditions more than the level of interest rates would predict.

relatively slow; while interest rates are at very low levels, it will be difficult to induce a sufficiently stimulative effect.

An expansionary policy beyond the low and stable rates may be preferable at the zero lower bound. A practical limitation in the current situation is brought about by the zero floor on the nominal interest rate, but this constraint on monetary policy operating through the interest rate channel need not be irremovable. In fact, the zero bound stems from the choice of governments and central banks to set the interest rate on coin, currency and the reserves of commercial banks with the central bank to zero. To illustrate this clearly, we adopt the presentation by Buiter and Panigirtzoglou (2003) in the following.[15] Let i denote the instantaneous nominal interest rate on non-monetary government debt, i_c is the interest rate on currency, i_R the nominal interest rate on commercial bank reserves with the central bank, γ the instantaneous marginal carry cost on bonds, γ_c the instantaneous marginal carry cost of currency and γ_R the instantaneous marginal carry cost on commercial bank reserves with the central bank. Then, considering the superior liquidity of base money, the following floor on the nominal interest rate on bonds holds:

$$i \geq Max\{i_c + \gamma - \gamma_c, i_R + \gamma - \gamma_R\} \tag{2.10}$$

If the carry cost terms γ, γ_c and γ_R are omitted, and base money is treated as a homogenous aggregate rather than a sum of currency and commercial bank reserves at the central bank, there is a single interest rate on base money i_M. Accordingly, we set $i_M = i_C = i_R$ and assume $\gamma = \gamma_C = \gamma_R = 0$. Equation (2.10) is then written as

$$i \geq i_M \tag{2.11}$$

The zero bound on nominal rates is no longer operative when a negative interest rate on base money is possible. An identical conclusion is obtained if a carry tax on base money is imposed in line with the suggestion by Goodfriend (2000), instead of a negative interest rate. These proposals are therefore considered interchangeably in the following paragraphs. Administrative problems for such measures arise from the fact that coin and currency are bearer bonds; as the holders are anonymous and ownership is sufficient for the bonds to be payable, it is difficult to make the bondholders actually pay the coupon for the issuer. The BOJ could avoid this problem by making currency subject to an expiration date and a conversion process during which

[15] The authors called the policy of paying a negative interest rate on base money "Gesell money", after Silvio Gesell, the perhaps best-known proponent of taxing currency.

the actual taxing of currency takes place. This can take the form of attaching coupons or stamps to the currency. After a certain date, say 1st January 2006, only the yen stamped/issued/added with a coupon on and after that date by the BOJ would be legal tender. The other component of monetary base, reserves at the central bank, pose no problem as the ownership is well known.

Freedman (2000) expressed concern about the amount of currency in circulation abroad, in case the US economy was to fall to a liquidity trap and a tax on currency to be adopted. In the case of Japan however, the magnitude of currency abroad is not likely to be a case of major concern. Recent data available from the Bank for International Settlements (2005) suggest that the share of cross-border foreign currency positions (assets) of banks vis-à-vis non-banks, denominated in yen, amounted to only 5.9% of all such positions in September 2004. This contrasts with a share of 59.5% in US dollars or 21.5% in euro.[16] Of course, currency substitution away from the yen toward dollars and euro could become significant if a currency tax is imposed, even if little evidence of significant substitution effects from yen in favour of foreign currencies has previously been found. Bryant (2000) noted that if substitution toward foreign currencies took place, it would increase pressure on home currency depreciation - something only welcome in the case of deflation in the home economy. Moreover, even if some (informal) cash transactions between economic agents could be handled in foreign currencies, legal measures could be imposed to ensure the settlement of all formal cash transactions in yen.

Note that our results of the interest rate channel remaining potent even at very low or negative inflation rates straightforwardly encourage the adoption of a tax on currency in Japan's current situation. The problem would be more complex if we had actually found a structural break in the monetary transmission mechanism, in line with the previously-mentioned suggestions of a weakened transmission mechanism in the Japanese economy (e.g. Miyao, 2000; Kimura *et al.*, 2002; Fujiki and Shiratsuka, 2002; Okina and Shiratsuka, 2003). In simulations by Hunt and Laxton (2001), output variability and the probability of the economy entering a deflationary spiral increase as the average inflation target drops below 2 percent. If (in contrast to our results) the interest rate channel had sufficiently weakened in Japan, then a tax on currency would have been most effective if adopted well before the zero floor was reached, with the inflation rate still in sufficiently positive territory.

[16] Similarly, Ito (2005) argued that the anchor, invoice, and settlement currency roles of the yen in East Asia have not lived up to expectations.

Similarly to our results, Ahearne *et al.* (2002) argued on the basis of model simulations that monetary policy would have been effective to counteract deflationary pressures in the early 1990s. Given the non-performing loans problem and instabilities in the financial system, suggestions about continued monetary policy effectiveness - even with very slow impacts on economic activity and prices - may appear striking. However, Woo (1999) showed that the most undercapitalized banks were actually expanding their lending most rapidly in the early 1990s. This result probably occurs due to a lack of supervision and a weak regulatory environment in the Japanese financial system. Even so, due to regulatory pressure and closer market scrutiny, banks with better capital positions were increasing their lending more than undercapitalized ones in 1997.[17]

Interestingly, the finding that broad money has a positive impact on the price level is closely linked to the benefits of adopting the tax on currency. Traditional monetary theory tells us that the money-holding sector demands base money since it is a medium of exchange that, unlike short-term government bills that are used in the open market operations of the central bank, provides transaction services to its holders. Of course, at zero interest rates, there is no incentive to hold anything else but base money - the economy is in the traditional liquidity trap. But as the tax on currency lowers the interest rate floor, the incentive to hold assets included in broader monetary aggregates is revived. According to our results this provides another way of increasing the price level than the interest rate channel of the central bank. Note also that an increase in broad monetary aggregates, if successfully implemented, can also be seen as a credibility measure on the part of the Bank of Japan - possibly more credible than a simple inflation or a price level target alone.[18]

[17]Krugman (1998) and Posen (1998) pointed out the moral hazard problem in the Japanese context; banks with low or negative net worth have an incentive to lend too much, rather than too little, to high-risk, high-return projects. Positive outcome would accrue to the benefit of the management and owners, while they only have the small or nonexistent net worth to lose. Similarly, Posen (1998) argued that lending was readily available from the bursting of the bubble until mid-1997; the decline in investment until this date stemmed from excess capacity of firms and low net worth, while there was little evidence of a credit crunch. Meredith (1998) also mentioned analysis by the IMF suggesting that credit conditions had returned to levels consistent with historical relationships after the bursting of the asset price bubble in 1992. Finally, Motonishi and Yoshikawa (1999) claimed that a fall in corporate investment in 1992-1994 was caused by real rather than financial factors.

[18]Blanchard (2000) mentioned the credibility aspect in the context of an increase in high-powered money, coupled with a credible commitment not to reverse the monetary expansion in the future.

2.4.4 An Optimizing IS Equation

The previous analysis established the potential of the interest rate channel to stimulate output in the Japanese economy which has moved from positive inflation rates to the deflationary era. Yet, inflation and price level targeting benefit from a certain degree of forward-lookingness on the part of economic agents. Even if it is also perfectly feasible in a backward-looking framework, inflation targeting serves as a useful anchor for expectations if the central bank's inflation target is credible; this increases the chances of the central bank actually hitting its future target. Svensson (2000) showed how in a system with forward-looking aggregate demand and supply, inflation is more self-stabilizing under inflation targeting. The importance of forward-lookingness in the case of price level targets can be argued to be even more important, since under this approach inflation shocks are not treated as bygones and no base drift of the price level is allowed. Then, any lower-than-expected inflation rate *ex post* must be met by a higher inflation rate in the future if the price level target is to be successfully met. As argued in the methodology section, the real *ex ante* rate then works as an automatic stabilizer: a stimulating impact on output after a period of low inflation can be obtained even without a movement in the actual nominal interest rate. This is of special interest when interest rates are already close to or at the zero bound. Hence, it is of importance to actually obtain an estimate for the significance of the real *ex ante* interest rate in the Japanese context.

In order to determine the degree of forward-lookingness in the determination of output and the importance of the real *ex ante* interest rate, we estimated an expanded Euler equation for Japan of the form:

$$y_t = a_0 + a_1 y_{t-1} + a_2 y_{t-2} + \mu E_{t-1} y_{t+1} + \beta E_{t-1}[1/\kappa \sum_{j=0}^{\kappa-1} (i_{t+j} - \pi_{t+j+1})] + \eta_t \quad (2.12)$$

This specification is identical to the one used by Fuhrer and Rudebusch (2004) in their estimation of the Euler equation for output in the US. Importantly for our purposes, potentially longer-term interest rates can be used; these are determined by the value κ obtains in the specification. A credible commitment by the central bank to keep interest rates low for a long period of time would likely show up in low expected nominal interest rates, in addition to increasing the future rate of inflation. As noted by Kuttner and Posen (2004), the gains from such a commitment could become significant when policy is constrained by the zero lower bound. Krugman (1998) was among the first to suggest that an increase in the expected inflation rate and the correspondingly lower real interest rates represent the channels through which the Japanese monetary authorities must operate, as nominal interest rates have hit the zero bound. In the model by Eggertsson and Woodford (2003), the entire future paths of short-term real and nominal rates or very long term real rates matter for aggregate demand.

In this specification, we define inflation as $\pi = 400(lnP_t - lnP_{t-1})$, where P_t denotes the price level. The nominal interest rate is the Bank of Japan call rate, and the output gap is obtained from the OECD. The determination of potential output in an economy that has gone through a significant economic slump, such as Japan, is notoriously difficult. This is aggravated by the fact that some explanations for Japan's long recession are based on structural arguments (see Hayashi and Prescott, 2000). However, using a top-down approach such as the Hodrick-Prescott filter to obtain an estimate of potential output is arguably a worse alternative, since such an approach assumes average deviations of output to be zero over the entire period and therefore may seriously underestimate the actual output shortfall.[19] As a robustness test, we also report the results using an output gap that was obtained by extracting the trend with a Hodrick-Prescott filter.

The IS equation was estimated by GMM for the period 1980Q1-2003Q4, with the main results presented in Table 2 below. Similarly to Fuhrer and Rudebusch (2004), the set of instruments contains the lagged "endogenous" variables, i.e. lags 1-4 of the output, interest and inflation rate series; these were used to instrument for the expectations terms (actual outcomes) of future periods. In Table 2, results using both the OECD measure for potential output and the one obtained by the Hodrick-Prescott filter are presented. Moreover, longer-term interest rates are taken into account by considering different values of κ. A time series weighting matrix was used in the estimation, making the estimates robust to heteroskedasticity and autocorrelation of an unknown form.

[19]The Hodrick-Prescott filter is referred to here as a top-down approach, as in Kuttner and Posen (2004), for such a method filters macroeconomic time-series data based on limited assumptions. This is in contrast to bottom-up methods based on assumptions about the production function, for example.

Specification		Coefficient estimates and standard errors					
Output	κ	α_1	α_2	μ	β	Adj. R^2	p-value
OECD	1	0.409	0.078	0.520	-0.022	0.917	0.42
		(0.060)	(0.037)	(0.047)	(0.008)		
OECD	2	0.413	0.074	0.522	-0.022	0.916	0.39
		(0.060)	(0.037)	(0.047)	(0.009)		
OECD	3	0.446	0.072	0.488	-0.023	0.916	0.39
		(0.061)	(0.037)	(0.049)	(0.009)		
OECD	4	0.443	0.058	0.504	-0.024	0.917	0.39
		(0.061)	(0.037)	(0.049)	(0.009)		
OECD	5	0.443	0.063	0.500	-0.025	0.916	0.39
		(0.061)	(0.037)	(0.050)	(0.009)		
HP	1	0.463	0.085	0.466	-0.032	0.790	0.49
		(0.059)	(0.048)	(0.043)	(0.009)		
HP	2	0.472	0.081	0.466	-0.034	0.788	0.44
		(0.059)	(0.052)	(0.043)	(0.009)		
HP	3	0.494	0.074	0.443	-0.032	0.784	0.42
		(0.057)	(0.047)	(0.046)	(0.010)		
HP	4	0.504	0.056	0.452	-0.033	0.787	0.45
		(0.058)	(0.047)	(0.046)	(0.009)		
HP	5	0.504	0.059	0.448	-0.034	0.785	0.44
		(0.058)	(0.047)	(0.048)	(0.010)		

Table 2. GMM estimates of the structural IS equation

Note: Standard errors in parentheses. The model specification columns indicate the procedure used to calculate the trend output, together with the duration of the real interest rate in the forward-looking specification. Instruments for all specifications included lags 1-4 of inflation, interest rate and output series. Displayed p-value refers to a test for over-identifying restrictions.

The results from the GMM estimation are rather consistent across the different specifications. The coefficient on the first lag of the dependent variable is highly significant in all the specifications, whereas the coefficient on the second lag is not always significantly different from zero. The estimates for the coefficient on expected output are rather high and statistically significant, with a magnitude close to 0.50. Importantly for our case, the role of expectations is also found to be significant in the case of the real interest rate. The estimated coefficient for expected output is slightly lower in magnitude than the one estimated by Fuhrer and Rudebusch (2004) for the US; however, our estimates for the coefficient on the real *ex ante* interest rate are higher in most specifications and obtain a higher statistical significance. The validity of the over-identifying restrictions in our estimations (we have more instruments than parameters to estimate) was never rejected at conventional levels of significance.

The fact that we obtain the expected sign and significance on the coefficient of the real *ex ante* interest rate does not ensure that the policy actually implemented by the Bank of Japan at this time would have corresponded to optimal policy as described in the theoretical literature. In this regard, Iwamura *et al.* (2005) claimed that a commitment to monetary easing until prespecified conditions regarding the inflation rate are satisfied lacks history dependence. Moreover, the commitment of the central bank did not sufficiently influence market expectations about the future course of monetary policy during 1999-2004, as indicated by the non-negative spread between the actual real interest rate and its natural rate counterpart. Despite such claims about the features of actual policy, our finding of the significance of longer-term real interest rates is encouraging, especially in view of apparently major changes in the BOJ's approach towards deflation. Specifically, unconventional measures such as quantitative easing were suggested to be ineffective in the past and there seemed to be a willingness to lift interest rates at any first sign of recovery - the zero interest rate policy was temporarily abandoned in August 2000 in the middle of the deflation era. Such rhetoric possibly caused deflationary expectations to become more entrenched in the economy. In contrast, a positive change in the credibility of the policy commitment should further increase the significance of the real interest rate (in the short and long-term) for the determination of output. Finally and importantly, the significance of the real *ex ante* interest rate is an important prerequisite for the successful implementation of a price level targeting strategy.

2.5 Conclusions

The aim of our paper was to examine whether inflation or price level targeting would be feasible monetary policy strategies for Japan. The pursuit of either approach critically hinges on two main prerequisites: the interest rate must remain a potent tool for an economy undergoing disinflation and deflation, and the central bank must possess an instrument to influence the future price level. Using evidence from vector autoregressions and structural IS equations, we found that both conditions are satisfied for Japan, even if monetary transmission to economic activity and prices appears to be strikingly slow. Yet, the ongoing potency of the interest rate channel facilitates a commitment policy to keep rates low into the future even after deflationary pressure has dissipated. Expansion in broad monetary aggregates was found to be causal for both prices and real output in Japan. Parameter constancy of the reduced form VAR estimates suggests that a possible policy to induce additional stimulus *via* the interest rate channel would be the implementation of a tax on currency; this would make it possible to realize any desired negative interest rate. Estimations from the structural IS equations further provided evidence of the importance of expectations in the Japanese economy both in terms of the real *ex ante* interest rate and real output. The management of expectations plays a crucial role especially in the successful implementation of a price level targeting rule.

It is interesting to note that the suggestions for Japan to adopt an inflation or a price level target come at a rather different time than for other economies that have introduced inflation targeting. Usually, such a policy was initiated to control runaway inflation, or after a successful disinflationary process with no threat of deflation. The apparent unwillingness of the BOJ to adopt an inflation target may have originated from the belief that no instruments were available to influence future prices - deeds have been thought to count more than words in the credible management of expectations. Another frequently invoked rationale for inflation targeting, increased central bank independence, may be currently undesirable for the BOJ; our results show that fiscal measures, such as a tax on currency, may need to be implemented in unison with monetary policy in order to overcome the deflation problem.

There are some reasons why the price level targeting approach (as opposed to inflation targeting) would be especially suitable for Japan. Very low positive inflation rates and deflation would necessitate a significantly positive inflation rate in the future under a price level targeting approach, as any deflationary shocks would need to be reversed. Under such a regime, any disinflationary shock is tackled immediately, with a greater likelihood of ensuring that the persistent deflationary era does not reappear. Moreover, with nominal rates still close to zero, a smaller variability in policy rates would lower the probability of the zero floor becoming binding again. If a price level target is eventually (significantly) overshot, contractionary monetary policy of some degree may need to be implemented, with potential costs for the economy. This, however, has to be weighed against the alternative of Japan returning to stagnant growth and deflation. Finally, neither inflation nor price level targeting would compromise the BOJ's ultimate goal of price stability that supports medium to long-term sustainable growth.

Bibliography

[1] Ahearne, A., Gagnon, J., Haltimaier, J., Kamin S., 2002. Preventing Deflation: Lessons from Japan's Experience in the 1990s. Board of Governors of the Federal Reserve System, International Finance Discussion Paper no. 729.

[2] Amisano, G., Giannini, C., 1997. Topics in Structural VAR Econometrics. Springer-Verlag, Berlin.

[3] Arai, Y., Hoshi, T., 2004. Monetary Policy in the Great Recession. In: Hutchison, M., Westermann, F. (Eds.), Economic Stagnation in Japan, CesIFO MIT Conference Volume.

[4] Bank for International Settlements, 2005. BIS Quarterly Review, March 2005.

[5] Bank of Japan, 2000. On Price Stability. Bank of Japan Publications.

[6] Bayoumi, T., 2001. The Morning After: Explaining the Slowdown in Japanese Growth in the 1990s. Journal of International Economics 53, 241-259.

[7] Benhabib, J., Schmitt-Grohé, S., Uribe, M., 2001. The Perils of Taylor Rules. Journal of Economic Theory 96, 40-69.

[8] Benkwitz, A., Lütkepohl, H., Wolters, J., 2001. Comparison of Bootstrap Confidence Intervals for Impulse Responses of German Monetary Systems, Macroeconomic Dynamics 5, 81-100.

[9] Berg, C., Jonung, L., 1999. Pioneering Price Level Targeting: The Swedish Experience 1931-1937. Journal of Monetary Economics 43, 525-551.

[10] Blanchard, O., 2000. Bubbles, Liquidity Traps, and Monetary Policy. In: Mikitani, R., Posen, A. S. (Eds.), Japan's Financial Crisis and Its Parallels to US Experience. Institute for International Economics, Washington DC.

[11] Breitung, J., Brüggemann, R., Lütkepohl, H., 2004. Structural Vector Autoregressive Modeling and Impulse Responses. In: Lütkepohl, H., Krätzig, M. (Eds.), Applied Time Series Econometrics. Cambridge University Press, Cambridge.

[12] Bryant, R. C., 2000. Comment on Overcoming the Zero Bound on Interest Rate Policy. Journal of Money, Credit and Banking 32 (4) Part 2, 1036-1050.

[13] Buiter, W. H., Panigirtzoglou, N., 2003. Overcoming the Zero Bound on Nominal Interest Rates with Negative Interest on Currency: Gesell's Solution. The Economic Journal 113, 723-746.

[14] Candelon, B., Lütkepohl, H., 2001. On the Reliability of Chow-Type Tests for Parameter Constancy in Multivariate Dynamic Models. Economics Letters 73, 155-160.

[15] Coenen, G., Wieland, V., 2003. The Zero-Interest-Rate Bound and the Role of the Exchange Rate for Monetary Policy in Japan. Journal of Monetary Economics 50 (5), 1071-1101.

[16] Coenen, G., Wieland, V., 2004. Exchange-Rate Policy and the Zero Bound on Nominal Interest Rates. AEA Papers and Proceedings 94 (2), 80-84.

[17] Dolado, J. J., Lütkepohl, H., 1996. Making Wald Tests Work for Cointegrated VAR Systems. Econometric Reviews 15, 369-386.

[18] ECB, 2003. The ECB's Monetary Policy Strategy. ECB Press Release, 8 May 2003.

[19] Eggertsson, G., Woodford, M., 2003. The Zero Bound on Interest Rates and Optimal Monetary Policy. Brookings Papers on Economic Activity 2003 (1), 139-211.

[20] Fischer, S., 1994. Modern Central Banking. In: Capie, F., Goodhart, C., Fischer, S., Schnadt, S. (Eds.), The Future of Central Banking. Cambridge University Press, Cambridge.

[21] Freedman, C., 2000. Comment on Overcoming the Zero Bound on Interest Rate Policy. Journal of Money, Credit and Banking 32 (4) Part 2, 1051-1057.

[22] Fuhrer, J. C., Rudebusch, G. D., 2004. Estimating the Euler Equation for Output. Journal of Monetary Economics 51, 1133-1153.

[23] Fujii, E., 2004. Exchange Rate Pass-Through in the Deflationary Japan: How Effective is the Yen's Depreciation for Fighting Deflation? CESifo Working Paper no. 1134.

[24] Fujiki, H., Shiratsuka, S., 2002. Policy Duration Effect under the Zero Interest Rate Policy in 1999-2000: Evidence from Japan's Money Market Data. Bank of Japan Monetary and Economic Studies, January 2002.

[25] Gaspar, V., Smets, F., 2000. Price Level Stability: Some Issues. National Institute Economic Review 174, 68-79.

[26] Goodfriend, M., 2000. Overcoming the Zero Bound on Interest Rate Policy. Journal of Money, Credit and Banking 32 (4), 1007-1035.

[27] Hayashi, F., Prescott, E. C., 2000. The 1990s in Japan: A Lost Decade. Minneapolis Federal Reserve Working Paper no. 607, November.

[28] Hunt, B., Laxton, D., 2001. The Zero Interest Rate Floor (ZIF) and Its Implications for Monetary Policy in Japan. IMF Working Paper no. 01/186.

[29] Ito, T., 2004. Inflation Targeting and Japan: Why Has the Bank of Japan Not Adopted Inflation Targeting? NBER Working Paper no. 10818.

[30] Ito, T., 2005. The Exchange Rate in the Japanese Economy: The Past, Puzzles, and Prospects. The Japanese Economic Review 56 (1), 1-38.

[31] Ito, T., Mishkin, F. S., 2004. Two Decades of Japanese Monetary Policy and the Deflation Problem. NBER Working Paper no. 10878.

[32] Iwamura, M., Kudo, T., Watanabe, T., 2005. Monetary and Fiscal Policy in a Liquidity Trap: The Japanese Experience 1999-2004. NBER Working Paper no. 11151.

[33] Jang, K., Ogaki, M., 2003. The Effects of Japanese Monetary Policy Shocks on Exchange Rates: A Structural Vector Correction Model Approach. Bank of Japan Monetary and Economic Studies, February 2003.

[34] JMulTi, 2004. Analysing Multiple Time Series with JAVA. Chair of Econometrics, Humboldt University Berlin. Available at www.jmulti.com.

[35] Johansen, S., 1995. Likelihood Based Inference in Cointegrated Vector Autoregressive Models. Oxford University Press, Oxford.

[36] Jung, T., Teranishi, Y., Watanabe, T., 2001. Zero Bound on Nominal Interest Rates and Optimal Monetary Policy. Kyoto Institute of Economic Research Discussion Paper no. 525.

[37] Kasa, K., Popper, H., 1997. Monetary Policy in Japan: A Structural VAR Analysis. Journal of the Japanese and International Economies 11, 275-295.

[38] Kashyap, A., 2002. Sorting Out Japan's Financial Crisis. NBER Working Paper no. 9384.

[39] Kato, R., Ui, T., Watanabe, T., 1999. Asymmetric Effects of Monetary Policy: Japanese Experience in the 1990s. Bank of Japan Working Paper no. 99-2, Research and Statistics Department.

[40] Kim, S., 1999. Do Monetary Policy Shocks Matter in the G-7 Countries? Using Common Identifying Assumptions about Monetary Policy across Countries. Journal of International Economics 48, 387-412.

[41] Kimura, T., Kobayashi, H., Muranaga, J., Ugai, H., 2002. The Effect of the Increase in Monetary Base on Japan's Economy at Zero Interest Rates: An Empirical Analysis. IMES Discussion Paper no. 2002-E-22.

[42] Krugman, P., 1998. It's Baaack: Japan's Slump and the Return of the Liquidity Trap. Brookings Papers on Economic Activity 1998 (2), 137-205.

[43] Kuttner, K. N., Posen, A. S., 2004. The Difficulty of Discerning What's Too Tight: Taylor Rules and Japanese Monetary Policy. North American Journal of Economics and Finance 15, 53-74.

[44] Lanne, M., Lütkepohl, H., Saikkonen, P., 2002. Comparison of Unit Root Tests for Time Series with Level Shifts. Journal of Time Series Analysis 23 (6), 667-685.

[45] Lipworth, G., Meredith, G., 1998. Background Paper: A Reexamination of Indicators of Monetary and Financial Conditions. In: Aghevli, B. B., Bayoumi, T., Meredith, G. (Eds.), Structural Change in Japan: Macroeconomic Impact and Policy Challenges. International Monetary Fund, Washington DC.

[46] Lütkepohl, H., 1991. Introduction to Multiple Time Series Analysis. Springer-Verlag, Berlin.

[47] Lütkepohl, H., 2004. Vector Autoregressive and Vector Error Correction Models. In: Lütkepohl, H., Krätzig, M. (Eds.), Applied Time Series Econometrics. Cambridge University Press, Cambridge.

[48] McCallum, B. T., 1988. Robustness Properties of a Rule for Monetary Policy. Carnegie-Rochester Conference Series on Public Policy 29 (Autumn), 173-203.

[49] McCallum, B. T., 2003. Japanese Monetary Policy, 1991-2001. Federal Reserve Bank of Richmond Economic Quarterly 89 (1) Winter 2003, 1-31.

[50] Meredith, G., 1998. Monetary Policy: Summary of Staff Views. In: Aghevli, B. B., Bayoumi, T., Meredith, G. (Eds.), Structural Change in Japan: Macroeconomic Impact and Policy Challenges. International Monetary Fund, Washington DC.

[51] Miyao, R., 2000. The Role of Monetary Policy in Japan: A Break in the 1990s? Journal of the Japanese and International Economies 14, 366-384.

[52] Miyao, R., 2002. The Effects of Monetary Policy in Japan. Journal of Money, Credit, and Banking 34 (2), 376-392.

[53] Miayo, R., 2005. Use of the Money Supply in the Conduct of Japan's Monetary Policy: Re-Examining the Time-Series Evidence. The Japanese Economic Review 56 (2), 165-187.

[54] Mishkin, F. S., 1999. Comment on Policy Rules for Inflation Targeting. In: Taylor, J. B. (Ed.), Monetary Policy Rules. University of Chicago Press, Chicago.

[55] Morsink, J., Bayoumi, T., 2000. Monetary Policy Transmission in Japan. In: Bayoumi, T., Collyns, C. (Eds.), Post-Bubble Blues: How Japan Responded to Asset Price Collapse. International Monetary Fund, Washington DC.

[56] Motonishi, T., Yoshikawa, H., 1999. Causes of the Long Stagnation of Japan during the 1990s: Financial or Real? Journal of the Japanese and International Economies 13, 181-200.

[57] Okina, K., 1993. Market Operations in Japan: Theory and Evidence. In: Singleton, K. J. (Ed.), Japanese Monetary Policy. University of Chicago Press, Chicago.

[58] Okina, K., Shiratsuka, S., 2003. Policy Commitment and Expectation Formations: Japan's Experience under Zero Interest Rates. IMES Discussion Paper no. 2003-E-5.

[59] Pantula, S. G., 1989. Testing for Unit Roots in Time Series Data. Econometric Theory 5, 256-271.

[60] Ploberger, W., Krämer, W., Kontrus, K., 1989. A New Test for Structural Stability in the Linear Regression Model. Journal of Econometrics 40, 307-318.

[61] Posen, A. S., 1998. Restoring Japan's Economic Growth. Institute for International Economics, Washington DC.

[62] Reifschneider, D., Williams, J. C., 1999. Three Lessons for Monetary Policy in a Low Inflation Era. Finance and Economics Discussion Paper no. 1999-44. Federal Reserve Board, Washington DC.

[63] Rotemberg, J. J., Woodford, M., 1999. Interest-Rate Rules in an Estimated Sticky-Price Model. In: Taylor, J. B. (Ed.), Monetary Policy Rules. University of Chicago Press, Chicago.

[64] Rudebusch, G. D., Svensson, L. E. O., 1999. Policy Rules for Inflation Targeting. In: Taylor, J. B. (Ed.), Monetary Policy Rules. University of Chicago Press, Chicago.

[65] Saikkonen, P., Lütkepohl, H., 2000a. Testing for the Cointegrating Rank of a VAR Process with Structural Shifts. Journal of Business & Economic Statistics 18, 451-464.

[66] Saikkonen, P., Lütkepohl, H., 2000b. Trend Adjustment Prior to Testing for the Cointegrating Rank of a Vector Autoregressive Process. Journal of Time Series Analysis 21, 435-456.

[67] Shioji, E., 2000. Identifying Monetary Policy Shocks in Japan. Journal of the Japanese and International Economics 14, 22-42.

[68] Sims, C. A., 1992. Interpreting the Macroeconomic Time Series Facts: The Effects of Monetary Policy. European Economic Review 36, 975-1000.

[69] Svensson, L. E. O., 1999a. How Should Monetary Policy Be Conducted in an Era of Price Stability? In: New Challenges for Monetary Policy, a Symposium Sponsored by the Federal Reserve Bank of Kansas City, 26-28 August 1999, 195-259.

[70] Svensson, L. E. O., 1999b. Price Level Targeting vs. Inflation Targeting: A Free Lunch? Journal of Money, Credit and Banking 31 (3) part 1, 277-295.

[71] Svensson, L. E. O., 2000. Open-Economy Inflation Targeting. Journal of International Economics 50 (1), 155-183.

[72] Svensson, L. E. O., 2001. The Zero Bound in an Open Economy: A Foolproof Way of Escaping from a Liquidity Trap. Bank of Japan Monetary and Economic Studies, Special Edition February 2001, 277-312.

[73] Svensson, L. E. O., 2003. Escaping from a Liquidity Trap and Deflation: The Foolproof Way and Others. Journal of Economic Perspectives 17 (4), 145-166.

[74] Teräsvirta, T., 1994. Specification, Estimation, and Evaluation of Smooth Transition Autoregressive Models. Journal of the American Statistical Association 89, 208-218.

[75] Teräsvirta, T., 2004. Smooth Transition Regression Modeling. In: Lütkepohl, H., Krätzig, M. (Eds.), Applied Time Series Econometrics. Cambridge University Press, Cambridge.

[76] Towe, C., 1998. Structural Change and Information Content of Financial Variables for Monetary Policy. In: Aghevli, B. B., Bayoumi, T., Meredith, G. (Eds.), Structural Change in Japan: Macroeconomic Impact and Policy Challenges. International Monetary Fund, Washington DC.

[77] Ueda, K., 1993. A Comparative Perspective on Japanese Monetary Policy: Short-Run Monetary Control and the Transmission Mechanism. In: Singleton, K. J. (Ed.), Japanese Monetary Policy. University of Chicago Press, Chicago.

[78] Warne, A., 2005. Structural VAR - version 0.32. Available at http://www.texlips.net/svar/.

[79] West, K. D., 1993. An Aggregate Demand-Aggregate Supply Analysis. In: Singleton, K. J. (Ed.), Japanese Monetary Policy. University of Chicago Press, Chicago.

[80] Williams, J. C., 1999. Simple Rules for Monetary Policy. Finance and Economics Discussion Paper Series no. 1999-12. Federal Reserve Board.

[81] Woo, D., 1999. In Search of "Capital Crunch": Supply Factors Behind the Credit Slowdown in Japan. IMF Working Paper no. 99/3.

[82] Woodford, M., 2003. Interest and Prices: Foundations of a Theory of Monetary Policy. Princeton University Press, Princeton.

2.6 Appendices

Appendix A

Data Sources

The following variables were obtained from the OECD Main Economic Indicators Database, vol. 2004, release 10: Gross domestic product (volume), the 3-month interest rate on the certificate of deposit, monetary aggregate M2+CDs, monetary aggregate broadly defined liquidity, monetary aggregate M1, the consumer price index and the GDP deflator.

The output gap was obtained from the OECD Economic Outlook Database, Vol. 75.

The call money rate was obtained from the IFS Database, series 15860B..ZF...

The series for monetary base and banknotes in circulation were obtained from the Bank of Japan long-term time series database.

All the series are at a quarterly frequency. The series on gross domestic product, banknotes in circulation, broadly defined liquidity, M2+CDs, M1, and the monetary base were seasonally adjusted, as reported by the OECD and the Bank of Japan, while the consumer price index was seasonally adjusted using a Census X-11 procedure by the author. All nominal monetary aggregates were deflated by the consumer price index in order to obtain the components of money stock in real terms.

Appendix B
Unit Root and Cointegration Tests

Augmented Dickey-Fuller Test

Series	Det. term	Lagged differences	Test stat.
$\Delta^2 y$	c	5 (AIC,HQ,SC)	-2.63**
Δy	c	2 (AIC,HQ,SC)	-3.16**
y	c, t	0 (AIC,HQ,SC)	-1.10
$\Delta^2 p$	c	5 (AIC, HQ)	-1.83
$\Delta^2 p$	c	4 (SC)	-1.54
Δp	c	2 (AIC,HQ,SC)	-2.67*
p	c, t	3 (AIC,HQ)	-0.37
p	c, t	0 (SC)	0.68
$\Delta^2 i$	c	7 (AIC,HQ)	-3.41***
$\Delta^2 i$	c	3 (SC)	-3.78***
Δi	c	2 (AIC,HQ)	-3.46***
Δi	c	0 (SC)	-5.89***
i	c, t	3 (AIC,HQ,SC)	-2.80

* indicates significance at 10% level, ** at 5% and *** at 1% level.
The order specification criteria in parentheses: AIC=Akaike, HQ=Hannan-Quinn, SC=Schwarz-criteria.
c and t denote constant and trend as deterministic terms, respectively.
All series except the interest rate in logarithms.
Maximum lag order set at 10, sample 1982Q1-2003Q4.

Saikkonen-Lütkepohl Cointegration Test

Series	Det. term	no. of lags	Coint. rank	test statistic
			0	47.53***
y, p, i	$c, S97Q2$	2 (AIC)	1	11.79*
			2	5.22**
			0	89.71***
y, p, i	$c, S97Q2$	1 (HQ,SC)	1	23.77***
			2	8.30***
			0	43.60***
y, p, i	$c, t, S97Q2$	2 (AIC)	1	2.31
			2	0.00
			0	54.13***
y, p, i	$c, t, S97Q2$	1 (HQ,SC)	1	11.57
			2	0.15

* indicates significance at 10%, ** at 5% and *** at 1% level.
c and t denote constant and trend as deterministic terms, respectively.
The order specification criteria in parentheses: AIC=Akaike, HQ=Hannan-Quinn, SC=Schwarz-criteria.
Prefix S denotes date of shift dummy.

Appendix C
Misspecification Tests

Benchmark model:

Q^*_{16}	132.88 [0.15]
LMF_5, LMF_4, LMF_1	1.09 [0.34], 1.02 [0.45], 1.44 [0.18]
$LJB(s_3^2), LJB(s_4^2)$	5.55 [0.14], 12.95 [0.00]
$ARCH_{LM}(16)$(eqs. 1, 2, 3)	8.22 [0.94] 10.57 [0.84] 23.26 [0.11]

Note: p-values in brackets.
Q* denotes the adjusted Portmanteau test statistic for autocorrelation.
LMF 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.

Model with M2+CDs:

Q^*_{16}	152.00 [0.11]
LM_5, LM_4, LM_1	47.21 [0.38], 37.10 [0.42], 7.03 [0.63]
$LJB(s_3^2), LJB(s_4^2)$	7.50 [0.06] 4.06 [0.26]
$ARCH_{LM}(16)$(eqs. 1, 2, 3)	9.06 [0.91] 7.85 [0.95] 14.58 [0.56]

Note: p-values in brackets.
Q* denotes the adjusted Portmanteau test statistic for autocorrelation.
LM is the Lagrange multiplier type 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.

Appendix D
Stability Tests

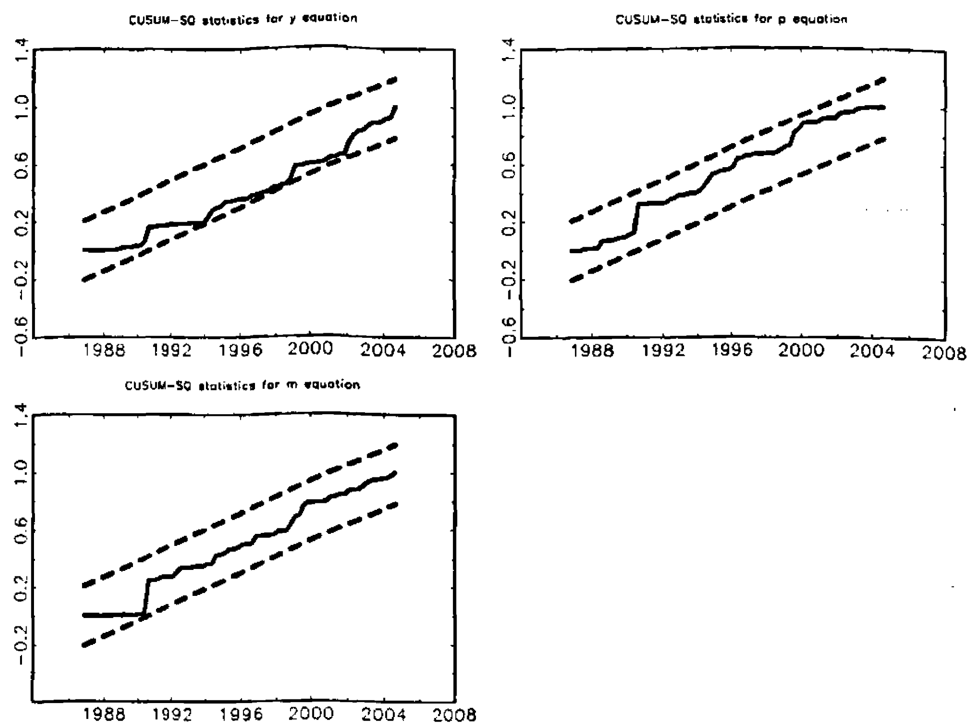
break date	chow_ss	boot p-val	chow_fc	boot p-val	break date	chow_ss	boot p-val	chow_fc	boot p-val
1993 Q1	100.6221	0.0045	0.924	0.7395	1995 Q1	70.7218	0.221	0.7344	0.955
1993 Q4	76.8267	0.109	0.7722	0.9265	1995 Q2	69.4657	0.2435	0.7214	0.9565
1994 Q3	75.4849	0.1365	0.7161	0.961	1995 Q3	68.0252	0.278	0.6889	0.973
1995 Q2	69.4657	0.2435	0.7214	0.9595	1995 Q4	70.1573	0.247	0.7214	0.9545
1996 Q1	71.2398	0.24	0.7426	0.935	1996 Q1	71.2398	0.2185	0.7426	0.947
1996 Q4	56.2095	0.4945	0.8514	0.804	1996 Q2	71.0463	0.226	0.7826	0.9065
1997 Q3	77.3222	0.1085	0.9862	0.661	1996 Q3	70.0936	0.14	0.8206	0.848
1998 Q2	54.9614	0.64	0.8638	0.845	1996 Q4	56.2095	0.4605	0.8514	0.807
1999 Q1	44.3914	0.824	0.7123	0.944	1997 Q1	63.8514	0.2795	0.8947	0.729
1999 Q4	NaN	NaN	0.7041	0.942	1997 Q2	66.2643	0.236	0.9283	0.684
2000 Q3	NaN	NaN	0.8516	0.746	1997 Q3	77.3222	0.1125	0.9862	0.6595
2001 Q2	NaN	NaN	1.0406	0.481	1997 Q4	68.1951	0.277	0.9486	0.7245
2002 Q1	NaN	NaN	0.5294	0.963					
2002 Q4	NaN	NaN	0.3377	0.989					
2003 Q3	NaN	NaN	0.421	0.856					

Chow sample split (chow_ss) and forecast (chow_fc) test statistics for benchmark model.
Bootstrapped p-values based on 2,000 replications

equation	test statistic	p-value
y_t	6.265	0.755
p_t	13.170	0.170
i_t	3.823	0.890

Recursive fluctuation tests for benchmark model.

Note: p-values based on 4,999 bootstrapping replications.



CUSUM of squares test, model with money, 5% significance level.

break date	chow_ss	boot p-val	chow_fc	boot p-val	break date	chow_ss	boot p-val	chow_fc	boot p-val
1993 Q1	17.798	0.400	0.670	0.703	1995 Q1	20.290	0.258	0.619	0.850
1993 Q4	24.417	0.111	0.616	0.832	1995 Q2	18.898	0.320	0.645	0.817
1994 Q3	20.831	0.227	0.570	0.910	1995 Q3	17.657	0.393	0.630	0.842
1995 Q2	18.898	0.302	0.645	0.823	1995 Q4	20.234	0.255	0.640	0.824
1996 Q1	19.890	0.260	0.651	0.811	1996 Q1	19.890	0.265	0.651	0.814
1996 Q4	21.729	0.183	0.741	0.632	1996 Q2	19.870	0.258	0.680	0.775
1997 Q3	19.051	0.327	0.819	0.578	1996 Q3	21.624	0.214	0.700	0.717
1998 Q2	7.548	0.882	0.835	0.547	1996 Q4	21.729	0.159	0.741	0.627
1999 Q1	-17.717	0.994	0.618	0.901	1997 Q1	21.113	0.206	0.752	0.620
1999 Q4	NaN	NaN	0.642	0.863	1997 Q2	21.248	0.193	0.771	0.586
2000 Q3	NaN	NaN	0.739	0.714	1997 Q3	19.051	0.322	0.819	0.595
2001 Q2	NaN	NaN	0.897	0.442	1997 Q4	13.073	0.646	0.848	0.529
2002 Q1	NaN	NaN	0.689	0.731					
2002 Q4	NaN	NaN	0.618	0.745					
2003 Q3	NaN	NaN	1.160	0.257					

Chow sample split (chow_ss) and forecast (chow_fc) test statistics for model with M2+CDs. Bootstrapped p-values based on 2,000 replications

Appendix E
Linearity Tests

Hypothesis	Transition variable i_{t-1}	i_{t-2}	i_{t-3}
Output equation			
H_0	0.94	0.94	0.80
H_{04}	0.70	0.58	0.44
H_{03}	0.70	0.77	0.62
H_{02}	0.87	0.95	0.86
Price equation			
H_0	0.47	0.28	0.30
H_{04}	0.28	0.35	0.34
H_{03}	0.66	0.24	0.61
H_{02}	0.36	0.33	0.13
Interest rate equation			
H_0	0.63	0.27	0.73
H_{04}	0.64	0.30	0.68
H_{03}	0.65	0.31	0.47
H_{02}	0.28	0.27	0.57

Table: p-values of linearity tests of benchmark VAR-model

Note: tests are based on the regression: $y_t = \beta_0' z_t + \sum_{j=1}^{3} \beta_j' \tilde{\mathbf{z}}_t s_t^j + u_t^*$, $t = 1, ..., T$.

where $u_t^* = u_t + R_3(\gamma, c, s_t)\theta' \mathbf{z}_t$ with the remainder $R_3(\gamma, c, s_t)$.

Here $\mathbf{z}_t = (\mathbf{w}_t', \mathbf{x}_t')'$ is a vector of explanatory variables, $\mathbf{w}_t' = (1, y_{t-1}, ..., y_{t-p})'$, and $\mathbf{x}_t = (x_{1t}, ..., x_{kt})'$ is a vector of exogenous variables. s_t is a continuous transition variable, γ is a slope parameter, and $c = (c_1, ..., c_K)'$ is a a vector of location parameters. $\theta = (\theta_0, \theta_1, ..., \theta_m)'$ is a $((m+1)\times 1)$ parameter vector and $u_t \sim iid(0, \sigma^2)$.

The hypotheses are as follows:

$H_0 : \beta_1 = \beta_2 = \beta_3 = 0$

$H_{04} : \beta_3 = 0$

$H_{03} : \beta_2 = 0 \mid \beta_3 = 0$

$H_{02} : \beta_1 = 0 \mid \beta_2 = \beta_3 = 0$

For details on the test, see Teräsvirta (1994, 2004).

Chapter 3

Japanese Broadly Defined Liquidity during Disinflation

3.1 Introduction

Changes in Japan's economic environment have created a formidable challenge for the conduct of Japanese monetary policy. The strong growth and positive inflation rates of the 1980s were followed by a financial bubble and eventually by deflation in the second half of the 1990s. Similarly, the importance of monetary aggregates in the policy framework of the Bank of Japan (BOJ) has undergone severe changes. The role of broad money M2+CDs diminished due to observed instability during the financial bubble of the early 1990s. This coincides with reduced interest in monetary aggregates in other major economies - the monetary targeting pursued by the German Bundesbank and the subsequent role of the monetary pillar in the policy framework of the European Central Bank can be regarded as major exceptions. The policy importance of the narrowest aggregates was revived in 2001 when current account balances of the commercial banks at the BOJ were chosen as the operating target of policy. In contrast, the behaviour and demand for broad monetary aggregates especially in the disinflationary and deflationary era have received less attention.

This study sets out to examine the demand for broadly defined liquidity in Japan during the years 1981-2004. The aim of the paper and our motivations are threefold. Firstly, a study of this broad monetary aggregate can be considered interesting in its own right - we are aware of only one study (at least in English!) that has discussed the relationship between broadly defined liquidity and economic activity in a vector error correction framework. More specifically, a paper by the Bank of Japan (2003) examined whether a long-run equilibrium relationship exists between broad money (both M2+CDs and broadly defined liquidity) and economic activity. Secondly, we are interested in the demand for a broad monetary aggregate during the disinflationary period that finally led to deflation in 1995. The puzzling observation of recent years is that the increases in monetary base implied by the policy of quantitative easing have not been reflected in portfolio adjustment to broader monetary aggregates. Even with limited controllability of broad money by the central bank, money stock broadly defined may still serve as an information variable for subsequent developments in real output and, more importantly, prices. In this regard, the relative absence of studies concerning broad money demand during deflation is rather striking. Thirdly, a recently renewed interest in broad monetary aggregates stems from their ability to yield valuable information on possibly destabilizing asset price developments. The latter may have an impact on the central bank's ultimate objective of price stability. This is especially interesting in the case of Japan, where the bursting of the asset price bubble in the early 1990s had profound implications for developments in the real economy.

In line with the consensus view in the money demand literature after the cointegration revolution, the framework adopted is a vector error correction (VEC) model that allows for the inclusion of both long-term cointegration and short-term relationships. We estimate both nominal and real money demand functions in the reduced VEC form and examine the dynamics and stability of the estimated real liquidity demand model. Imposing a structural break during the bursting of the financial bubble, the use of I(2) techniques is completely avoided. The latter methodology is often required to investigate money demand if the money stock and price series need to be differenced twice in order to obtain stationary time series for econometric inference. We find that a relatively stable and economically meaningful demand system can be established for the Japanese broadly defined liquidity during 1981-2004. This finding arises despite the asset price bubble, instabilities in the financial system and the onset of deflation in the Japanese economy during our observation period. These findings are in contrast to some of the earlier literature which suggests that the relationship between real economy and the broad money stock has become unstable in the recent years. We confirm the significance of share prices for Japanese broad money demand, while little evidence of important currency substitution effects is detected. Finally, the investigation of system dynamics is based on impulse response analysis in the context of a structural vector error correction (SVEC) model for the Japanese broadly defined liquidity.

Our analysis should not be considered as advocating anything resembling a monetary targeting strategy for the Bank of Japan. Such an approach would hardly be feasible with a very broad aggregate that even includes government bonds in its construction. Rather, we are particularly interested in broad money for its information content; this requires a stable relationship with the variables of interest. The Bank of Japan (2003) has claimed the information value of money for policy to be higher when the long-run relationship between the real money stock and the real economy or between the money stock and the price level is stable, and the causality from money to economic activity is strong. We provide evidence of stable long-run relationships between real liquidity, output and share prices, together with causality from nominal liquidity to both consumer and share prices, to be satisfied for our investigated aggregate. The latter finding is of interest especially in the context of the persistent deflation problem in the Japanese economy.

This paper is structured as follows. The next section presents the definitions of Japanese broad money stock and discusses some of the previous literature concerned with the demand for Japanese broad money. Section 3 elaborates on the theoretical issues pertinent to our research question. The fourth section presents our empirical results, including a statistical analysis of the time series properties of the series, and estimation of the reduced form VEC models. An analysis of the dynamic relationships between the variables is tackled in a separate section, in the context of a structural VEC model. The final section is a conclusion of our findings.

3.2 Definitions of Japanese Broad Money Stock and Previous Literature

In this section, we present the definitions of two indicators of Japanese broad money stock and comment on the role of the monetary aggregates in the conduct of monetary policy by the Bank of Japan. Some of the previous studies concerning the demand for Japanese broad money are also discussed.

The Bank of Japan currently compiles four indices for money stock: M1, M2+CDs, M3+CDs and broadly defined liquidity. The most prominent of the broad monetary aggregates, both in the policy framework and as a source of research interest, has historically been M2+CDs. Since the first declaration of its importance (M2) as a leading indicator for prices in July 1975, money stock has been considered an important information variable. In

July 1978, the BOJ started producing projections for the growth rate of M2 (later M2+CDs) over the coming quarter ("quarterly foresights"), even if money stock never officially obtained the role of an intermediate target. On one hand, the role of money in the monetary policy framework regained its prominence with the adoption of a regime of quantitative easing in 2001, even if only in the form of narrow money. On the other hand, the BOJ considers the emergence and bursting of the asset price bubble, and the resulting non-performing loans problem, to have influenced the relationship between money stock and economic activity (Bank of Japan, 2003). Such arguments can be regarded as demoting the perceived information value of money for policy.[1] Our study provides evidence that is in stark contrast with the claim of a reduced information content of broad money stock.

In this section, we focus the discussion on M2+CDs and broadly defined liquidity. Their definitions are described below in Table 1. Money holders in Japan comprise corporations, individuals, local public authorities, municipal enterprises and public corporations. Excluded are money issuers of M2+CDs, the central government and financial institutions involved in loan business (Bank of Japan, 2004).

Aggregate	Composition
$M2 + CDs$	M1 (cash currency + demand deposits) + time deposits + certificates of deposit
Broadly defined liquidity	M2+CDs + deposits (including CDs) held with post offices + agricultural cooperatives + money trusts and loan trusts of domestically licensed banks (excluding foreign trust banks) + government bonds (including financing and treasury bills) + bonds with repurchase agreement + bank debentures + investment trusts + money deposited other than money in trust and foreign bonds

Table 1. Composition of two indicators of Japanese broad money stock: M2+CDs and broadly defined liquidity.

[1] Nevertheless, the BOJ closely followed monetary developments even after the bursting of the asset price bubble. As an example, statements accompanying interest rate cuts in 1995 mentioned that sluggishness in money growth has influenced interest rate policy (Towe, 1998).

Interestingly, a comparison between Japanese broadly defined liquidity and the M3 aggregate used in the context of the so-called monetary pillar of the European Central Bank reveals some similarities. In particular, money market fund shares/units, repurchase agreements and securities lendings with cash collateral are included in broadly defined liquidity but not in the narrower indicators of Japanese money stock (Bank of Japan, 2004). The Eurozone M3 includes M2 plus repurchase agreements, money market fund shares/units and money market paper, together with debt securities up to two years. Furthermore, Japanese broadly defined liquidity includes instruments that could be considered appropriate to be contained in M2+CDs; however, they have been omitted from the latter due to limited information from data sources (only outstanding amounts at the end of period are available). It is also worth noting that our interest in Japanese broad money is mainly based on the persistent deflationary environment in Japan and therefore on the link between broadly defined liquidity and other macroeconomic variables, including the price level. Gerlach (2004) operationalized the ECB's two-pillar monetary policy framework by suggesting a model where the low-frequency inflation component depends on past monetary growth, consistently with the ECB view that the link between money growth and inflation would be most apparent in the medium to long run.

It is plausible that broadly defined liquidity could be more stable than M2+CDs simply due to its broad coverage. More specifically, the broad liquidity variable includes investment trusts and bank debentures that were affected by portfolio adjustment to M2+CDs (especially to highly liquid bank deposits), which took place in the wake of financial system anxieties in 1997. In this regard, the Bank of Japan (2003) has argued that problems in the financial system caused a break in the stable relationship between M2+CDs and economic activity after 1997, making it difficult to obtain evidence about current and future developments in economic activity and prices, and more generally about the effects of changes in the monetary stock. In Figure 1 below, we depict the development (annual growth rates) of the two broad monetary aggregates in real terms, broadly defined liquidity and M2+CDs, from 1981 to 2004. The figure confirms the slightly more stable movement of broadly defined liquidity, even if the two aggregates track each other quite closely.

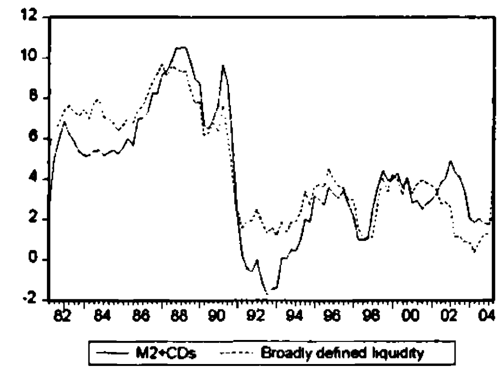


Figure 1. Annual growth rates of M2+CDs and broadly defined liquidity (in real terms).

As mentioned earlier, the prominence of M2+CDs in the monetary policy framework has been a major reason for the dominance of this aggregate in studies of money demand in Japan. Moreover, data on broadly defined liquidity has only been published from 1980 onwards. In the following, we discuss some of the previous literature about the demand for broad Japanese monetary aggregates.

To our knowledge, only one study, conducted by the Bank of Japan (2003), has examined the relationship between broadly defined liquidity and economic activity in a vector error correction framework. The estimated trivariate system included real money stock, real GDP, and an interest rate spread between a 10-year government bond and a 3-month CD rate.[2] In contrast to the investigated system with M2+CDs, a long-run equilibrium relationship including broadly defined liquidity was detected even after end-1997 when anxieties over the financial system emerged. The aggregate was also found to lead developments in both nominal and real GDP. However, the adjustment coefficient on broadly defined liquidity in the VECM was actually found to be positive, suggesting no convergence to a long-run equilibrium and the variable was therefore judged not to be economically meaningful.

Inspired by the role of broad money supply (M2+CDs) as an information variable, Miyao (2005) presented time series evidence with recent data

[2]The yield on a 10-year government bond may not be an optimal rival interest rate for broadly defined liquidity, as government bonds are included in this broad indicator, and the variable should by construction appear rather stable to changes of interest rates among the components of money.

suggesting that the linkage between broad money and income or prices had disappeared in the 1990s. More specifically, the author could not reject the absence of a stable long-run relationship between money, output and prices. Moreover, neither time deposits nor bank loans were useful in predicting movements in output in the late 1990s. Using an earlier sample, Tsukuda and Miyakoshi (1998) claimed that causality from M2+CDs to income weakened or disappeared in the 1980s. Kimura (2001) suggested that, when accounting for the financial system shock in the estimated VEC model, the long-run relationship between M2+CDs and real income actually remained stable even when the years 1998 and 1999 were included in the analysis.

The measurement of wealth effects was given a prominent role in the study by Sekine (1998), which examined the demand for broad money M2+CDs during 1975-1994, a period of financial liberalization and the bubble economy. The money demand system accounted for wealth effects and financial liberalization and established a stable demand function. The wealth variable used was the System of National Accounts (SNA) definition of wealth. In addition to financial assets, this variable also covered non-financial assets such as land, housing and inventories. Due to the availability of the wealth series at an annual frequency only, it was interpolated to a quarterly series. The result of money demand stability over a period of the asset price bubble, financial innovation and deregulation was also confirmed in papers by Amano and Wirjanto (2000) and Bahmani-Oskooee (2001).

Kameda, Kyoso and Yoshida (1998) examined long-run relations between monetary aggregates and other macroeconomic variables in Japan. A stable long-run relationship between real M2+CDs and real GDP was detected during 1960-1996. It is of interest that the authors were unable to reject the I(2) property of the money stock series but the econometric analysis was conducted under the assumption of the series being integrated of order one. An earlier study by Yoshida and Rasche (1990) investigated the structure and stability of the demand for real broad money, again M2+CDs, in Japan during 1956-1989. The authors discovered a stable relationship between real broad money and income. Yoshida (1990), who also included stock price volatility in his model, similarly found a relatively stable real money demand function for M2+CDs during 1968-1989.

As is evident from the preceding discussion, empirical literature on the demand for broad monetary aggregates is rather limited for the disinflationary and deflationary period - only Miyao (2005) has presented evidence through 2003. The literature concerning the broadly defined liquidity aggregate is even sparser. We claim that by adequately accounting for the structural break during the time of the financial bubble, a relatively stable and economically meaningful broad money demand function can be obtained by using I(1) cointegration techniques only. The next section comments on theoretical issues pertinent to our research question.

3.3 Theoretical Considerations

In this section, we discuss some theoretical issues relevant to our research question. These include a long-run specification for money demand and a discussion of the rationale for the inclusion of the individual variables, with a special emphasis on the characteristics of broadly defined liquidity. In this context, the possible leading indicator or causal properties of broad money with respect to asset prices are commented on. Finally, we elaborate on the importance of the money stock (both narrowly and broadly defined) in the context of the current deflationary episode.

3.3.1 Long-Run Specification of Money Demand

As suggested by Ericsson (1999), most standard theories propose at least two reasons for money demand. These are the inventory function in order to smooth out differences between income and expenditure streams, and one of several assets included in a portfolio. This would lead to the characterization of long-run money demand in the form:

$$M^d/P = f(Y, \tilde{R}) \tag{3.1}$$

Here M^d represents nominal money demanded, P is the price level, Y is a scale variable and $\tilde{R}$ is a vector of returns on various assets. In applied work, a semi-log linear form is generally used to illustrate an empirical approximation to equation (3.1), as follows:

$$m_t^d - p_t = \gamma_0 + \gamma_1 y_t + \gamma_2 R_t^{own} + \gamma_3 R_t^{out} + \gamma_4 \pi_t \tag{3.2}$$

In this specification, m_t^d, p_t and y_t are expressed in logarithms; $\gamma_0, \gamma_1, \gamma_2, \gamma_3$ and γ_4 are coefficients; R_t^{own} and R_t^{out} represent the return to assets included and excluded in money, and π_t is the inflation rate. The expected signs for the coefficients are $\gamma_1 > 0, \gamma_2 > 0, \gamma_3 < 0$, and $\gamma_4 < 0$. Long run price homogeneity of money demand is generally assumed. Moreover, some theories explicitly predict certain values for the coefficients, such as $\gamma_1 = 0.5$ in the Baumol-Tobin model or $\gamma_1 = 1$ in the quantity theory of money. Often some measures to indicate the influence of wealth on long-run money demand are also included; alternatively, the absence of these is used to justify a value $\gamma_1 > 1$. However, the characteristics of broadly defined liquidity would motivate a somewhat different specification from (3.2). In particular, our estimated system for (both nominal and real) broadly defined liquidity omits interest rates on domestic monetary assets included and excluded in the examined aggregate R_t^{own} and R_t^{out}, but includes the (log real) Tokyo share price index $topix_t$. Moreover, the importance of external influences in the form of currency substitution CS_t is investigated by testing the significance of different proxies for such phenomena in the system. We therefore effectively consider

$$(m_t^d - p_t) = f(y_t, topix_t, \pi_t, CS_t) \tag{3.3}$$

Issues pertaining to the inclusion of the different variables are discussed in the following paragraphs, while their endogeneity/exogeneity properties are discussed in the context of model estimation.

Regarding the price index p_t, it has been commonplace to use the GDP deflator in the analysis of Japanese money demand functions (e.g. Bank of Japan, 2003; Kimura, 2001; Kimura *et al.*, 2002; Sekine, 1998) instead of the consumer price index. In the Japanese context of falling prices, however, the behaviour of the consumer price index is interesting, as it is widely used as a measure of the magnitude of the current deflation. Moreover, according to the most recent clarifications by the BOJ, the current zero rate policy framework is to stay in place until the CPI inflation rate (excluding fresh food) stays at or above zero percent for a few months and there is no forecast of the economy falling back into deflation by the BOJ board members (Ito and Mishkin, 2004). These considerations justify the use of the CPI both as a deflator and as a measure of inflation (defined as $\pi_t = p_t - p_{t-1}$) in our paper. An investigation of the dynamics between the nominal money stock and the CPI also motivates us to begin from a demand specification for nominal broad money where the price level is included as a separate endogenous variable.

An interesting feature of broadly defined liquidity is that it ought to be little affected by portfolio adjustment between different components of money. As stated by the Bank of Japan (2004), the aggregate is a large scale index that has a tendency towards stable transition even in the face of capital inflow and outflow of various financial instruments. Consequently, we do not regard interest rates on the components of money as very meaningful measures for broadly defined liquidity. Similarly, all the rates used to compute the (maximum) interest rate on rival assets for M2+CDs by Sekine (1998) are included in the broad liquidity aggregate. Our choice, then, is to use the inflation rate π_t as the opportunity cost variable for broadly defined liquidity. Such a variable is usually regarded as a measure of return for holding real assets instead of money. However, it may also be viewed as representing the kind of portfolio adjustment process of economic agents (see e.g. Lütkepohl and Wolters, 1999). Note that our estimated system includes two other types of variables that could theoretically assume the role of an opportunity cost indicator. These include the share price index and various indicators to capture currency substitution effects, even if the latter were omitted from the final model due to statistical insignificance. These variables are discussed next.

Considerable emphasis has been placed on the evolution of asset prices in Japan due to the financial bubble that led to a collapse of Japanese asset values, including share prices, in the early 1990s. This also motivates the inclusion of the (log) real Tokyo share price index $topix_t$ in our paper. Friedman (1988) suggested that the effect of share prices on money demand depends on whether the wealth or the substitution effect prevails. Regarding the first effect, a rise in share prices leads to an increase in nominal wealth and, assuming a wider fluctuation of share prices than income, a higher ratio of wealth to income. This can then lead to a higher money to income ratio or a lower velocity. Considering money as an asset in an investor's portfolio, higher real share prices may alternatively make equities more attractive in terms of their return, inducing a substitution effect. The share price variable would then assume the role of rival return for money. Interestingly, the European Central Bank has partly attributed the stubbornly high headline figures for M3 growth of the recent years to the persistent weakness of the stock markets and uncertainty in the financial markets, seemingly subscribing to the existence of a strong substitution rather than a wealth effect (see e.g. ECB, 2001).

Even if the wealth effect of equity prices prevailed, as suggested by our findings, we cannot consider the share price variable to be a comprehensive indicator of Japanese household wealth. Iwaisako (2003) showed that the proportion of equities in household financial wealth was decreasing throughout the 1990s. The author reported survey data by the Bank of Japan indicating a peak in the share of equities in total financial wealth at 10.6% in 1990, declining to 7.2% in 1999, close to the value of 6.1% in 1981.[3] Impor-

[3]However, Bayoumi (2001) pointed to the possibility of wealth effects explaining the growth slowdown in Japan, and suggested that these would primarily operate through

tantly, the inclusion of share prices should be seen in the light of examining the causal or leading indicator properties of broad money for asset prices. This hinges on its ability to yield valuable information on possibly destabilizing asset price developments that may also have an impact on the central bank's ultimate objective of price stability. This discussion has been especially prominent in the context of the monetary pillar, used by the European Central Bank, which focuses on the developments of broad money M3. Issing (2002) argued that the close monitoring of monetary and credit developments is beneficial in order to limit the emergence of unsustainable asset valuations. Emphasizing the information properties of money, Nelson (2003) suggested that money conveys information about monetary conditions that is not summarized by the short-term interest rate only. In particular, an inclusion of money in New-Keynesian macroeconomic models incorporates the entire spectrum of yields in the analysis that matter both for aggregate and money demand. As pointed out by Masuch *et al.* (2003), the importance placed by Nelson (2003) on a broader set of yields than those observed in the securities markets may provide answers to the association between episodes of money growth, the build-up of financial imbalances and asset price bubbles. During financial instability, a comparison between a short-term rate controlled by the central bank and a benchmark rate such as the one given by a Taylor rule may not correspond to the perception of current monetary conditions by market participants.[4]

The importance of a currency substitution term in a money demand system is relevant for policymaking, as the unpredictability of money demand may increase when such external effects are not properly accounted for. This could, at the extreme, reduce the effectiveness of monetary policy. Moreover, while the interest rate differential (defined in terms of domestic components of money) is not a very meaningful measure for broadly defined liquidity, profit opportunities abroad may still provide a rival rate of return for this aggregate. It is important to note, however, that foreign interest rates and exchange rates affect a part of broadly defined liquidity, as the aggregate includes foreign bonds, issued by non-residents in domestic or foreign markets held by money holders (Bank of Japan, 2004).[5] Their share in the total

share prices, as land prices had divergent effects on property owners and those with no land.

[4]In a similar vein, Morsink and Bayoumi (2000) found that the Japanese bubble in the late 1980s was driven by broad money and loan shocks rather than by interest rate shocks, and that interest rate policy was not surprisingly expansionary during that period.

[5]This includes samurai bonds (yen-denominated foreign bonds issued by non-residents in the domestic market), but excludes foreign bonds issued by domestic residents in foreign markets.

amount outstanding of broadly defined liquidity is small though; as an example, figures for July 2004 showed them to amount to 3.3% of the total only.

Following the approach of Tavlas (1996) for the Japanese M1, we experiment with various different indicators for currency substitution CS_t. These include two alternative measures of the exchange rate and an interest rate differential with respect to a US short-term interest rate. As the first exchange rate indicator, we include the rate of appreciation of the real effective exchange rate $\Delta reer = reer_t - reer_{t-1}$, where $reer$ denotes the (log) level of the real effective exchange rate, defined in terms of the foreign currency price of domestic currency. Ho *et al.* (2005) provided evidence that the Asian currencies have been recently trading in a more "effectively" oriented fashion rather than as a dollar bloc. This finding emphasizes the co-movements of the respective currencies vis-à-vis the currencies of the trading partners. Moreover, Tavlas (1996) suggested that the low demand by Japanese residents for foreign currency-denominated assets could stem from historically successful inflation control in this economy, motivating the use of the real rate. The canonical importance of the US-yen exchange rate inspires our second proxy for currency substitution, y^{US-JP}. It is derived from the idea of Corsetti and Pesenti (1999) that the expected GDP growth differentials between two currency zones could explain the movements of the (nominal) bilateral rate. In similar vein to the idea's implementation by Artis and Beyer (2004) for euro area money demand, we used the actual growth differentials between the US and Japan as a proxy for expected exchange rate appreciation. Finally, an interest rate indicator i^{US-JP} captures the difference between the US and Japanese 3-month CD rates. An increase in the return of foreign assets compared with the Japanese ones could induce portfolio adjustment favouring the former. However, such investment shifts would only result in currency substitution (defined in the context of our aggregate) if they concern foreign assets other than bonds, as the latter form part of broadly defined liquidity.

This subsection has discussed the rationale for the inclusion of the various variables in our demand specification for broadly defined liquidity, taking into consideration the special features of this broad aggregate. The next subsection elaborates on the usefulness of money stock (both narrowly and broadly defined) for a central bank operating close to the zero interest floor.

3.3.2 Money Stock and the Zero Lower Bound on Interest Rates

The control of money stock is of special interest for a central bank close to or at the zero bound, as conventional monetary stimulus through lower nominal interest rates may be unavailable. In particular, a Keynesian view of the liquidity trap taking hold as soon as nominal rates hit zero may be overly simplistic from a monetarist viewpoint. It is only where a composite asset of money and bills is a perfect substitute for all other assets that an economy would also be in a liquidity trap in a monetarist sense (see e.g. Meltzer, 1999). In general, a change in the monetary policy stance, such as an increase in the monetary base, causes a re-balancing of portfolios which is operative even at the zero lower bound through various channels (Kimura *et al.*, 2002). These include the relative asset-supply effect, where an increase in the monetary base may lead to changes in the relative prices of assets, a reduction of transaction costs due to a smaller probability of liquidity shortage, and the reduction of longer rates through the expectations channel.

McCallum (e.g. 1988, 2003) proposed a rule for controlling the monetary base, which can be written as

$$\Delta b_t = 5 - \Delta v_t^a + 0.5(5 - \Delta x_{t-1}) \tag{3.4}$$

Here b_t is the (log) monetary base, Δv_t^a is the average rate of base velocity growth over the past four years, and x_t is the (log) nominal GDP. The assumptions of a 2 percent inflation target and a 3 percent growth rate for real GDP yield a target value of 5 for nominal GDP growth. It is of interest that McCallum (2003) compared the actual values with those prescribed by the rule for Japan, finding that the Bank of Japan policy has been almost continuously too tight in terms of the monetary base since mid-1990.

In the event of instabilities in the money multiplier, even a successful following of such a rule does not guarantee a corresponding increase in broad money. This is evident in the context of the recent Japanese experience, where the high growth rates of the monetary base have not been reflected by corresponding increases in the broad money stock. However, stability and causality between economic activity and broad money could still justify the use of money broadly defined as an information variable for the monetary authority. Moreover, in the theoretical literature, an increase in broad monetary aggregates has been seen as a means of spurring expenditure even after the zero floor on interest rates has become binding and deflation has taken hold (see e.g. Goodfriend, 2000; Hetzel, 2003). Goodfriend (2000) distinguished between the liquidity services offered by narrow and broad money. Whereas transactions services offered by narrow money may satiate the public at a zero rate, liquidity services broadly defined still facilitate financial intermediation even at the interest rate floor. Assets included in broad money often serve as collateral for external finance and minimize the agents' exposure to the external finance premium. Then, the amount of broad money outstanding could be of interest in an economy that has hit the zero interest rate floor.

Conventional open market operations could prove powerless in a situation with very low interest rates. Indeed, when short-term interest rates are zero, commercial banks may keep the excess reserves resulting from operations with the central bank on their balance sheet. In such a situation, Goodfriend (2000) suggested that the central bank could purchase long-term bonds directly from the non-bank public. As the monetary base is expanded in exchange for long bonds, the marginal liquidity services yield on monetary assets declines. Through a portfolio adjustment in favour of less liquid assets, such as consumer durables and physical capital, asset prices are positively affected. This triggers a higher rate of consumption in relation to current income, with higher asset prices increasing investment. In sum, the portfolio rebalancing channel induces a positive impact on income, consumption and employment. Moreover, as monetary assets increase and asset prices rise, the external finance premium falls, having a positive effect on lending. Simultaneously, the balance sheets of both firms and households improve. Through this process working *via* the credit channel of monetary transmission, spending increases because the cost of borrowing has declined. Note that as our broad liquidity aggregate includes government bonds in the hands of the non-bank public, the amount of outstanding bonds included in this aggregate can be seen to illustrate the feasibility of such a policy. Government bonds included in the aggregate amounted to 85.0 trillion yen in June 2004, corresponding to 6.2% of broadly defined liquidity outstanding.

3.4 Empirical Evidence

This section presents the empirical evidence obtained in our analysis. We commence with a brief description of the data, followed by an investigation of its time series properties. Then, we present the estimation results from the reduced form VEC models. The dynamics of the model are examined in the context of impulse response analysis in a separate section.

3.4.1 About the Data

We use quarterly, seasonally adjusted data from the OECD Main Economic Indicators Database for the period 1981Q2-2004Q2. Our estimation sample covers a number of important phenomena: the build-up and bursting of the Japanese asset price bubble, disinflation that turned into deflation in 1995, financial system anxieties emerging in late 1997, and the start of the zero interest rate policy of the BOJ in 1999. The precise data sources are listed in Appendix A, including graphs of the series used in the estimation. All variables except the GDP growth and interest rate differentials between the US and Japan were transformed into logarithms. In the estimations, the software JMulTi (2004), version 3.11, was predominantly used. Exceptions are mentioned in the text.

3.4.2 Unit Root and Cointegration Tests

In order to determine the order of integration of the variables in our system, we employed augmented Dickey-Fuller (ADF) tests together with unit root tests with a structural break by Lanne *et al.* (2002). We followed the so-called Pantula principle (see Pantula, 1989) in the testing procedure, assuming a maximum integration order of two for all the series. ADF tests for the series in first differences, listed in Appendix B, suggested a rejection of a unit root for real output, consumer and stock prices and the real exchange rate, but a unit root could not be rejected for these variables in levels. The only exceptions are the interest and GDP growth rate differentials between the US and Japan, for which the levels of the variables were found to be stationary.[6]

Graphical observation of the series suggests that conventional unit root tests may prove inadequate so far as the monetary variables in our study are concerned. Instead, unit root tests with a structural break may be of particular relevance when the data includes a possible break point, such as the bursting of the asset price bubble that led to lower growth rates of the broad Japanese monetary aggregates. Evidence for our argument was already depicted in Figure 1 in terms of declining annual growth rates of broad money. The results from the unit root tests with a structural break for the real and nominal liquidity series are presented in Appendix B, with the break date

[6] At times we found it difficult to reject a unit root for the series in second differences even if a unit root was clearly rejected for the series in first differences. This finding arises when a large number of lags is used in the testing and could consequently be due to a power problem.

set at 1990Q3 (or Q1), consistently with the suggestion from a formal test. The results indicate that an adequate accounting of the break caused by the financial bubble leads to the rejection of a unit root for the monetary series in first differences. In contrast, ADF tests that do not account for the structural break in some cases suggested that a unit root could not be rejected for the series in first differences, consistent with a graphical examination of the series. In conclusion, we proceed under the assumption that the series for the monetary aggregates are integrated of order one when possible break points in the series are taken into account.[7]

Due to the nature of the underlying series, the cointegrating rank of the series was investigated using the tests proposed by Saikkonen and Lütkepohl (2000) that allow for the inclusion of level shifts. The tests proceed by estimating the deterministic terms by a GLS procedure, subtracting them from the observations and finally applying a Johansen type of test on the adjusted series.[8] We include two shift dummies in our tests, as well as in the subsequent estimation of the vector error correction system. These are, firstly, a shift dummy in the third quarter of 1990 at the time of the bursting of the financial bubble (denoted as $S90Q3$) and secondly, for 1997Q2 ($S97Q2$) coinciding with a consumption tax hike that led to a level shift in the consumer price index and notable output fluctuation, having been introduced during a modest recovery after the initial appearance of deflation.[9] Note that the former dummy captures a level shift in the first differences of the monetary series as the growth rates of broad liquidity decreased, but the second dummy is needed for the level shift in the consumer price index in levels. Graphical observation would support the inclusion of a deterministic trend in the cointegration testing and the tests were thus conducted with and without such a deterministic term. Furthermore, the tests were performed both in the nominal and real money demand systems; the latter are comprised of three variables only and are the primary source of statistical inference here (the

[7] If a break date in the consumer price index corresponding to the consumption tax hike of April 1997 is additionally taken into consideration in testing for unit roots for real liquidity, it would be necessary to include two break dates in the test. This could be achieved simply by a Saikkonen-Lütkepohl cointegration test on the liquidity series in first differences with two dummy variables. Such a test indeed indicates a stationary series, as a cointegration rank of zero was clearly rejected.

[8] The Johansen test (see Johansen, 1995) is a likelihood ratio test based on a reduced rank regression of a vector error correction specification.

[9] We have decided to leave another consumption tax hike - that of 1989Q2 - unaccounted for in the estimated model. The magnitude of the level shift of consumer prices is less apparent during the "inflation period", also with regard to its implications for output fluctuations during the booming economy.

currency substitution and the opportunity cost variables being treated in a separate context). We argue that our approach of focusing on three variables only is justified from the viewpoint that real money demand is the focus of the study; additionally, from an econometric perspective, the low power of cointegration tests with a high number of dimensions is somewhat alleviated when only three variables are being studied.

The results from the Saikkonen-Lütkepohl cointegration tests are somewhat dependent on the chosen lag length and the deterministic terms used. However, using a 5% level as a reliable indication of statistical significance, with constant and trend as the deterministic terms, most of the evidence from the systems including the real money variable suggests that the cointegration rank of zero can be rejected. Moreover, we are not able to reject a cointegrating rank of one. This also holds for the nominal money demand system when a constant is included as a deterministic term, although when a trend is also included a second cointegration relation cannot be rejected. In order to obtain more robust evidence for the real liquidity demand system, the sample was shortened by one year both from the beginning and from the end; again with a constant, trend and the shift dummy variables as the deterministic terms, a cointegrating rank of one could never be rejected in these alternative samples. Bivariate tests point to a cointegrating rank of either one or zero between real money and stock prices, a rank of one between output and stock prices, and a rank of one between real money and output. Given these results of the bivariate testing, it is not surprising to obtain (at a minimum) a cointegrating rank of one between the three variables. As a conclusion, we proceed to the estimation of the VEC model assuming that a cointegrating rank of one exists between the underlying variables. This is furthermore justified by our focus on the real liquidity demand system where all trivariate tests pointed to one cointegration relationship, while a second cointegration relation (additional to a money demand relationship) would be difficult to motivate.

3.4.3 Estimation of the Reduced Form VEC Models

A vector autoregressive model can be written as

$$x_t = A_1 x_{t-1} + \dots + A_p x_{t-p} + CD_t + u_t \tag{3.5}$$

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 $E(u_t) = 0$. The model given by (3.5) has a vector error correction representation as follows:

$$\Delta x_t = \Pi x_{t-1} + \Gamma_1 \Delta x_{t-1} + ... + \Gamma_{p-1} \Delta x_{t-p+1} + CD_t + u_t \tag{3.6}$$

When the variables are cointegrated, Π has reduced rank $r = rk(\Pi) < K$. It can be written as $\Pi = \alpha\beta'$, where α and β are $(K \times r)$ matrices that contain the loading coefficients and the cointegration vectors, respectively.

The reduced form VEC model was estimated for the period 1981Q2-2001Q2, with the vector of endogenous variables written as $x_t = (m_t, p_t, y_t, topix_t)'$. The lag order was chosen to be 3 on the basis of misspecification tests and considerations of an adequate number of lags to properly examine the monetary transmission mechanism. The observations 1981Q2-1982Q1 were used as presample values, leaving us with a total of T=89 observations. Restricting the two shift dummies and the trend in the cointegrating relation, we obtain the following long-run relationship with the Johansen ML procedure (with the standard errors in parentheses):

$$m_t = \underset{(0.167)}{0.868}\, p_t + \underset{(0.167)}{0.527}\, y_t + \underset{(0.015)}{0.161}\, topix_t - \underset{(0.018)}{0.013}\, S90Q3 - \underset{(0.012)}{0.000}\, S97Q2 + \underset{(0.001)}{0.007}t + ec_t \tag{3.7}$$

A normalization of the coefficient on broadly defined liquidity to one led to the presentation of the cointegration relation in the form of a possible money demand relationship. While the coefficients on the price level, output, share prices and the trend variable are all statistically significant, this is not the case for the two shift dummies at any conventional significance level.[10] Remember that a shift dummy on 1990Q3 was required in order to account for the structural break in the broad liquidity series, making it possible to characterize the series in first differences as stationary. As a consequence, we treat it as orthogonal to the cointegration relation, without restricting it to the long-run relationship. The shift dummy on 1997Q2 to account for the consumption tax increase was included in the short-run part of the model as an impulse dummy for the same quarter. This yields the following long-run relation:

$$m_t = \underset{(0.155)}{1.042}\, p_t + \underset{(0.155)}{0.425}\, y_t + \underset{(0.015)}{0.163}\, topix_t + \underset{(0.001)}{0.007}t + ec_t \tag{3.8}$$

[10]It is theoretically difficult to justify the inclusion of a trend in the cointegration relation and to argue that variables in our long-run money demand relation could be driven apart by a linear trend. However, the trending behaviour of the share price variable during the "boom and bust economy", when stock prices initially rapidly increased and then fell, could partly explain the statistical significance of the trend in the cointegration relation.

The biggest changes with respect to the long-run relationship represented by (3.7) are the somewhat lower coefficient on real output and the coefficient on prices that is very close to one. Testing the validity of long-run homogeneity of prices with respect to money amounts to setting the coefficient of prices to one in the β matrix of the VEC specification (3.6), which cannot be rejected at any conventional significance level. Again estimating with a lag length of 3, the Johansen methodology yields the following real liquidity demand relationship (with standard errors in parentheses):

$$(m-p)_t = \underset{(0.137)}{0.371}\ y_t + \underset{(0.016)}{0.170}\ topix_t \underset{(0.001)}{+0.008t} + ec_t \tag{3.9}$$

Note a coefficient on real income that is significantly different from unity. Interestingly, the coefficient is very close in magnitude to the value found in a money demand study by Sekine (1998) for M2+CDs - the author found a long-run income elasticity of 0.42 - suggesting that higher coefficients found in previous studies could reflect the omission of an important wealth variable. Indeed, the positive coefficient on the real share price variable seems to indicate that wealth effects, instead of substitution effects, prevail in the case of the equity price indicator. In a study concerning the demand for M2+CDs, Kimura (2001) found that the cointegrating relationship between money and income broke down after 1997Q3 if financial anxieties were not taken into account in the form of an additional variable in the system. We estimated the system until 1997Q3 and interestingly found very similar coefficients to our longer sample: 0.376 and 0.171 for real output and share prices, respectively, when the shift dummy for 1990Q3 was included in the cointegration relation. This finding could be interpreted to suggest that our broad aggregate seems relatively robust to the problems in the financial system that are said to have caused a breakdown in the relationship between money and economic activity.

We acknowledge the undesirable property of the maximum likelihood estimator in that it possibly produces distorted estimates in small samples (for details, see e.g. Brüggemann and Lütkepohl, 2004). We therefore conducted robustness tests for the cointegration parameters by shortening the sample by one year from the beginning and from the end of the estimation period. However, while the estimated β-coefficients of the cointegrating vectors show some variation to changes in the estimation sample, an alternative procedure, discussed in Lütkepohl (2004) by the name of the S2S procedure, did not produce markedly more robust results in our case, as shown in the table below. The estimates for the β-coefficients remain fairly robust also to estimating the long-run relation by the S2S procedure with a lag length of 4.

method	sample	lag length	β_y	SE(β_y)	β_{topix}	SE(β_{topix})
ML	1981Q2-2004Q2	3	0.371	0.137	0.170	0.016
ML	1982Q2-2004Q2	3	0.428	0.131	0.160	0.016
ML	1981Q2-2003Q2	3	0.355	0.144	0.170	0.017
S2S	1981Q2-2004Q2	3	0.430	0.136	0.162	0.016
S2S	1982Q2-2004Q2	3	0.505	0.129	0.149	0.015
S2S	1981Q2-2003Q2	3	0.419	0.142	0.161	0.017
S2S	1981Q2-2004Q2	4	0.341	0.152	0.173	0.018
S2S	1982Q2-2004Q2	4	0.405	0.139	0.161	0.017
S2S	1981Q2-2003Q2	4	0.336	0.156	0.169	0.018

Table 2. Estimated cointegration vectors with alternative methods and samples.

As we have dealt with the cointegration properties of the system already and a stationary variable may influence the cointegration relation above, we include our stationary opportunity cost variable, the inflation rate as an exogenous (unmodelled) variable in the system. Moreover, estimating the cointegration relation (3.9) by including the inflation rate yields a statistically insignificant coefficient with a wrong sign (0.042, with a standard error of 0.155). In accordance with the endogenous variables, the opportunity cost variable was included with 3 lags, together with its contemporaneous lag. A reduction to a subset model was subsequently achieved by a procedure whereby at each step the parameter with the lowest t-value of the entire system was checked and potentially eliminated from the model. A threshold value of 1.00 was specified; only variables with t-values higher than this threshold were then eventually maintained. Such an approach may be useful in a VAR/VEC system that typically includes many statistically insignificant coefficients. This yields the following equation for real broadly defined liquidity:

$$\begin{aligned}\Delta(m-p)_t \;=\; & \underset{(0.015)}{-0.068}[(m-p)_{t-1} \underset{(0.137)}{-0.371} y_{t-1} \underset{(0.016)}{-0.170} topix_{t-1} \underset{(0.001)}{-0.008} t_{t-1}] \quad (3.10)\\ & \underset{(0.090)}{+0.193}\Delta(m-p)_{t-1} \underset{(0.070)}{+0.164}\Delta(m-p)_{t-3} \underset{(0.036)}{-0.083}\Delta y_{t-3} \underset{(0.073)}{-0.936}\Delta p_t\\ & \underset{(0.110)}{+0.295}\Delta p_{t-1} \underset{(0.092)}{+0.205}\Delta p_{t-3} \underset{(0.001)}{-0.007} S90Q3 \underset{(0.004)}{+0.030} + u_{1t}\end{aligned}$$

Two issues are particularly noteworthy in the equation above. Firstly, both the possible convergence to equilibrium and its speed are of interest. Evidence on this is provided by the coefficient on the cointegration relation, considered to express the weight with which the long-run relation enters the respective equation. Despite the relatively low speed of convergence - the coefficient obtains a value of -0.068 - we obtain a negative sign for the coefficient, suggesting a stable model. This is in stark contrast with the finding by the Bank of Japan (2003) of the long-run relationship not being economically meaningful in the estimated system for broadly defined liquidity. Secondly, the coefficient on the current lag of the opportunity cost variable is statistically significant with the expected sign; moreover, the combined lags on the inflation rate obtain the expected negative sign in the long run.

The equations for real output and real share prices are written similarly as:

$$\begin{aligned}\Delta y_t &= \underset{(0.006)}{0.007}[(m-p)_{t-1} - \underset{(0.137)}{0.371} y_{t-1} - \underset{(0.016)}{0.170} topix_{t-1} - \underset{(0.001)}{0.008} t_{t-1}] \\ &\quad + \underset{(0.112)}{0.285}\Delta y_{t-1} + \underset{(0.010)}{0.022}\Delta topix_{t-2} + \underset{(0.106)}{0.218}\Delta y_{t-3} + \underset{(0.223)}{0.452}\Delta p_{t-1} \\ &\quad - \underset{(0.205)}{0.388}\Delta p_{t-3} - \underset{(0.008)}{0.015} I97Q2 + u_{2t}\end{aligned} \tag{3.11}$$

and

$$\begin{aligned}\Delta topix_t &= \underset{(0.431)}{2.408}[(m-p)_{t-1} - \underset{(0.137)}{0.371} y_{t-1} - \underset{(0.016)}{0.170} topix_{t-1} - \underset{(0.001)}{0.008} t_{t-1}] \\ &\quad + \underset{(0.094)}{0.565}\Delta topix_{t-1} + \underset{(0.833)}{0.975}\Delta y_{t-2} + \underset{(0.849)}{1.752}\Delta y_{t-3} + \underset{(0.095)}{0.307}\Delta topix_{t-3} \\ &\quad - \underset{(1.878)}{2.348}\Delta p_t - \underset{(1.672)}{1.991}\Delta p_{t-2} - \underset{(0.022)}{0.116} S90Q3 + \underset{(0.066)}{0.154}\, I97Q2 - \underset{(0.104)}{0.529} + u_{3t}\end{aligned} \tag{3.12}$$

As can be seen from the adjustment coefficients on the above equations, real output is clearly weakly exogenous in our system, whereas this is not the case for the real share price variable. The latter coefficient obtains a surprisingly high magnitude in absolute terms, even if the loading coefficients are dependent on the chosen normalization and therefore somewhat arbitrary, as discussed by Lütkepohl (2004). A high speed of adjustment in terms of the stock price variable could be the result of the "boom and bust economy" of our estimation sample and the correspondingly strong movements in the share price index at this time.

Is there any evidence for currency substitution in our demand system for broadly defined liquidity? We tested for the inclusion of the three various currency substitution variables one-by-one by estimating three separate systems. The subset model specification described above served as our starting point; the system was then re-estimated with the respective currency substitution variable. Finally, the coefficients on the currency substitution variable below the threshold level of 1.00 were also deleted from the system. This yields the coefficients on the various currency substitution indicators in the real liquidity equation of the VEC as follows (t-values in brackets):

System 1	Currency substitution variable: i^{US-JP}			
Lag	t	$t-1$	$t-2$	$t-3$
	—	—	-0.016 [-1.306]	—
System 2	Currency substitution variable: $\Delta reer$			
Lag	t	$t-1$	$t-2$	$t-3$
	—	—	0.008 [1.378]	—
System 3	Currency substitution variable: y^{US-JP}			
Lag	t	$t-1$	$t-2$	$t-3$
	—	-0.029 [-1.051]	—	0.050 [1.257]

Table 3. Coefficients on alternative currency substitution variables in real liquidity demand equation.
Note: the coefficients on i^{US-JP} and y^{US-JP} are multiplied by 100.

None of the coefficients on the different currency substitution variables in the equation for real liquidity are significant. Moreover, the limited significance of our currency substitution indicators was detected in the other equations of the VECM as well. However, the variable $\Delta reer$ obtained a positive coefficient (0.299 with a t-value of 2.035) on the third lag of the real share price equation. The difference between the GDP growth rates of US and Japan y^{US-JP} was significant, not surprisingly, in the real output equation at current and second lags. Due to the rather limited evidence of external influences in our system, we did not include a currency substitution variable in our final model, the three equations for which are then written as (3.10) - (3.12) above.

What could be the reason for the small importance of the currency substitution variable in the context of Japanese broad money? As claimed by Sekine (1998), the small or nonexistent significance of the exchange rate in the Japanese money demand functions could be partly due to historical regulations prohibiting the Japanese from maintaining overseas deposits. The amendment of the Foreign Exchange Law, effective from April 1998, may already have caused higher sensitivity of money demand to exchange rate movements but these changes are probably too recent to have a statistically significant impact in our sample. Moreover, as suggested by Tavlas (1996), there is little empirical evidence of considerable substitution between Japanese residents' holdings of non-yen denominated financial instruments and yen deposits. This data is obtainable from deposit figures for Shinkin banks that are comparable to savings and loans institutions in the US. In August 1994 - a date that falls roughly in the middle of our estimation sample - the amount of foreign currency deposits by residents amounted to only 1.1% of total deposits at these banks. As a final caveat, to the extent that changes in exchange and interest rates induce portfolio adjustment to foreign bonds within the broadly defined liquidity aggregate, no change would be observed in the level of broad liquidity outstanding.

Our subset model for real liquidity demand was submitted to various misspecification tests, as displayed in Appendix C. We performed the LM-tests for residual autocorrelation, and single-equation Jarque-Bera and ARCH-LM tests to detect nonnormality and ARCH-effects in the residuals, respectively. The Portmanteau test for autocorrelation was additionally implemented on the model without exogenous variables. The tests do not predominantly suggest a rejection of the estimated model, even if there is evidence of ARCH effects and nonnormality in the first equation of the VECM. However, the ARCH effects are only borderline significant at 5% level. Similarly, misspecification tests were conducted on the model for nominal liquidity demand. No evidence of autocorrelation or ARCH effects was found at a 5% significance level; only for the nominal liquidity equation of the VECM do we find evidence of nonnormality. Therefore, our models appear adequate for our analysis.

Recursive eigenvalue tests were conducted in order to examine the stability of the estimated system for real broadly defined liquidity, proposed for VAR models with cointegrated variables by Hansen and Johansen (1999). A concentrated likelihood function was used in the estimation, leading to a recursive estimation of the long-run part of the system only. Moreover, the confidence intervals for the eigenvalues were obtained on the basis of estimating the standard errors from the full sample. A significance level of 5% was used. The recursive eigenvalue analysis, together with other stability tests, is illustrated in Appendix D. The displayed recursive eigenvalue suggests that the bursting of the asset price bubble in the beginning of the 1990s does not go unnoticed in our system. However, after this time the diagnostic statistic does not indicate the existence of major instabilities, even if slight instability is detected toward the end of our estimation sample. Moreover, the values of the tau statistic are considerably smaller than the 5% critical value throughout the investigation period.

Further stability testing was performed by the Chow break point and sample split tests, where the null hypothesis of time invariant parameters was tested against the possibility of a change in the parameter values at a certain date. Our small sample results confirmed the finding by Candelon and Lütkepohl (2001) that the approximate χ^2 and F-distributions of the Chow test statistics are rather poor approximations, leading to high rejection rates. Accordingly, bootstrapped p-values were used, based on 1,000 replications. Our approach was to search through the maximum available sample instead of focusing on some particular individual break dates that could be of interest for the analysis. In both the Chow breakpoint and sample split tests, the p-values were always higher than the 5% critical value. Additionally, we conducted recursive fluctuation tests for the short-run parameters of the full unrestricted model (see e.g. Ploberger *et al.*, 1989).[11] Bootstrapped p-values for the empirical distributions were used, based on 4,999 replications. Again, these were never lower than the 5% level, suggesting that the parameters of the individual equations were relatively stable during our estimation period. We therefore concluded that the majority of our evidence points to a stable model.

We used our model for real broadly defined liquidity in order to formulate out-of-sample forecasts for real liquidity and real GDP 7 quarters ahead. The model was estimated until 2002Q3 and forecasts for 2002Q4-2004Q2 were compared with the actual outcomes for this period. Exogenous variables were not included in the models for the forecasting exercise. Additionally, forecasts for the nominal money stock and the price level were obtained using an identical methodology for the nominal liquidity model. These are displayed below in Figure 2, including the actual outcome (solid line), and the forecast (dotted line), together with the 95% confidence intervals. All the actual outcomes fall within the confidence intervals, even if our model is not able to capture remarkably well the rapid increase in export-led GDP growth that took place in the course of 2003. The forecast for the consumer price index is strikingly close to the actual outcome, although such a result also reflects limited fluctuations in Japanese consumer prices.

[11] This test was performed using the Structural VAR program by Warne (2005).

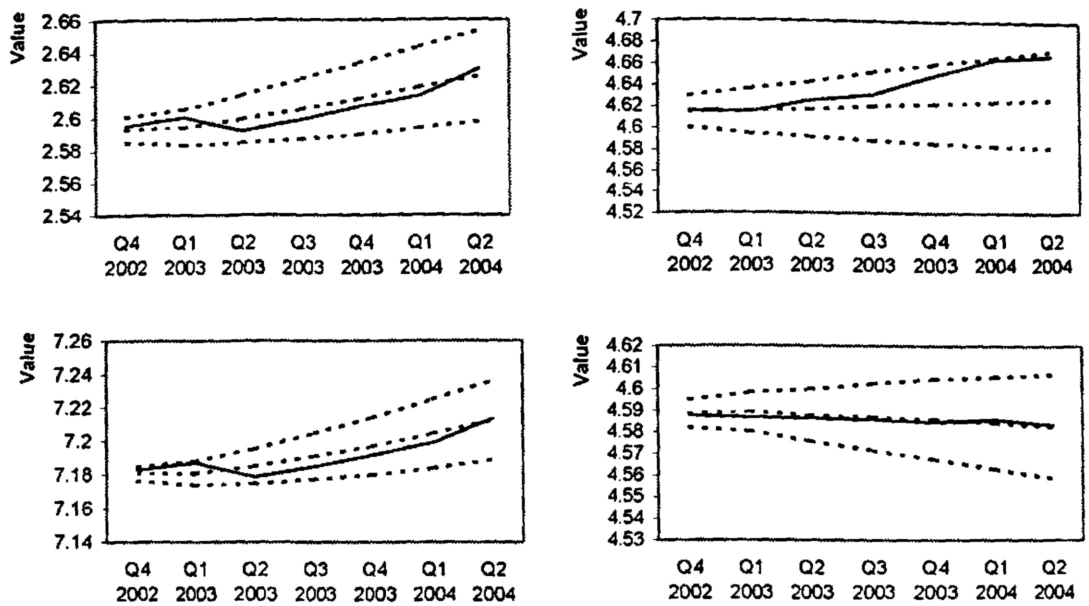


Figure 2. Forecasts for broadly defined liquidity (real, top-left; nominal, below-left), real GDP (top-right) and the consumer price index (below-right).

3.5 Impulse Response Analysis

The previous analysis established the existence of a relatively stable demand system for broadly defined liquidity in the Japanese economy during a time period when the economy experienced the financial bubble and its subsequent bursting, and disinflation that turned into deflation. However, little has been mentioned about the dynamics of our system, in addition to those pertaining to the cointegration relationship and in a limited fashion to the short-term part of the model. As it is difficult to draw an inference from the large number of coefficients of the model, the dynamics are discussed in this section in the context of impulse response analysis, where the effects of various shocks are traced through the system.

Impulse response analysis is conducted in our case using a structural vector error correction (SVEC) model. We consider two different identification schemes to construct the model for impulse response analysis. In the first case, identification is achieved by using short-run (contemporaneous) restrictions only, whereas the second set of identifying restrictions applies information from the previous cointegration analysis, together with possible considerations from economic theory. The first case is more common in the literature; a recursive ordering such as a Cholesky decomposition relies on a short-run timing scheme in order to identify the shocks. In the second case, identification obtains primarily through the use of long-run restrictions. The latter approach potentially avoids the problems of giving economic interpretations to reduced form VAR equations and to the shocks arising in the context of impulse response analysis. Of course, a recursive ordering is usually justified on the basis of economic considerations also. However, economic theory generally provides more suggestions about the long-run impacts of the shocks than about the mere contemporaneous effects. Finally, to our knowledge, only Jang and Ogaki (2003) have used long-run restrictions on a VECM system to conduct analysis on Japanese monetary policy. In contrast, none of the previously mentioned studies about Japanese broad money demand provide results from an impulse response analysis.

In order to derive the structural form, we focus on the effects of the fundamental shocks ε_t on the variables in the system x_t. These shocks are obtained from the structural form VECM (in what follows, we closely follow the presentation by Breitung *et al.*, 2004):

$$\mathrm{A}\Delta x_t = \Pi^* x_{t-1} + \Gamma_1^* \Delta x_{t-1} + ... + \Gamma_{p-1}^* \Delta x_{t-p+1} + \mathrm{B}\varepsilon_t \tag{3.13}$$

where $\varepsilon_t \sim (0, I_K)$. $\Pi^*$and $\Gamma_j^*(j = 1, ..., p-1)$ are structural form parameter matrices. The $(K \times K)$ matrix A is invertible and allows for the modelling of the instantaneous relations. Note that here the deterministic terms are omitted since they are not affected by the impulses hitting the system, nor do they affect such impulses themselves. The structural shocks, ε_t, are related to the model residuals by linear relations. Moreover, they are assumed to be mutually uncorrelated and therefore orthogonal. Because it is not possible to directly observe these structural shocks, certain assumptions are necessary for identification. In order to compute the responses to the structural shocks ε_t, the reduced form disturbances u_t have to be linked to them. When we premultiply (3.13) by A^{-1}, we obtain the reduced form (3.6), where $\Pi = \mathrm{A}^{-1}\Pi^*$and $\Gamma_j = \mathrm{A}^{-1}\Gamma_j^*(j = 1, ..., p-1)$. Specifying A to be an identity matrix, the reduced form disturbances are linked to the underlying structural shocks simply by

$$u_t = \mathrm{B}\varepsilon_t \tag{3.14}$$

It can be shown (see Breitung *et al.*, 2004) that the long-run effects of structural shocks ε_t can be written as

$$\Xi B \tag{3.15}$$

This matrix has rank $K-r$. Then, it can have at most r columns of zeros. The zero columns represent the r shocks with transitory effects (a zero long-run impact), whereas $k^* = K-r$ shocks have permanent effects. Because the matrix is of a reduced rank, each column of zeros stands for k^* independent restrictions. Exact identification of permanent shocks requires $k^*(k^*-1)/2$ restrictions, while $r(r-1)/2$ are needed for the transitory shocks. Together this yields $k^*r + k^*(k^*-1)/2 + r(r-1)/2$ restrictions - which is enough for a just-identification of the B matrix.

3.5.1 Identification with Contemporaneous Restrictions

In order to examine the dynamics between the nominal money stock and the price level, we return to the model for nominal broadly defined liquidity where the price level is included as an endogenous variable in the system. A cointegration relation for such a system was written in Equation (3.8) in Section 4 of the paper. Here, the identification of the model is achieved by the use of contemporaneous restrictions only, by means of a recursive identification scheme. Such an approach is very common in the literature, and is in fact effectively employed when a reduced VAR model is estimated and the errors are orthogonalized by means of a Cholesky decomposition. With a variable ordering of m_t, p_t, y_t, and $topix_t$, restrictions on the B matrix are written as

$$\begin{bmatrix} * & 0 & 0 & 0 \\ * & * & 0 & 0 \\ * & * & * & 0 \\ * & * & * & * \end{bmatrix} \tag{3.16}$$

where asterisks denote the unrestricted elements. Our set of contemporaneous restrictions implies that a shock to the nominal money stock may have an instantaneous impact on all the variables, whereas a shock to share prices does not have an instantaneous impact on any of the other variables. In our system, the equity price variable can therefore be considered the most responsive to changes in economic conditions. As a possible drawback, the outcome of the various shocks may be dependent on a particular ordering of the variables. Moreover, information on common stochastic trends is not explicitly taken into account in setting the restrictions.

Estimates for the contemporaneous impact matrix are obtained by maximum likelihood, starting from the reduced form estimates of the covariance matrix, subject to the restrictions imposed in the structural form. Numerical optimization methods in the form of a scoring algorithm are used (see Amisano and Giannini, 1997; Breitung *et al.*, 2004). Hall bootstrapped percentile confidence intervals at a 95% significance level were chosen to take into account parameter uncertainty. Apart from the satisfactory small sample properties of the bootstrap estimators in comparison with other asymptotic methodologies, this approach benefits from a built-in bias adjustment (see Benkwitz *et al.*, 2001). The number of bootstrap replications was set at 5,000. We depict in Figure 3 below the responses of the system variables to a shock in nominal money. Other impulse responses are depicted in Appendix E.

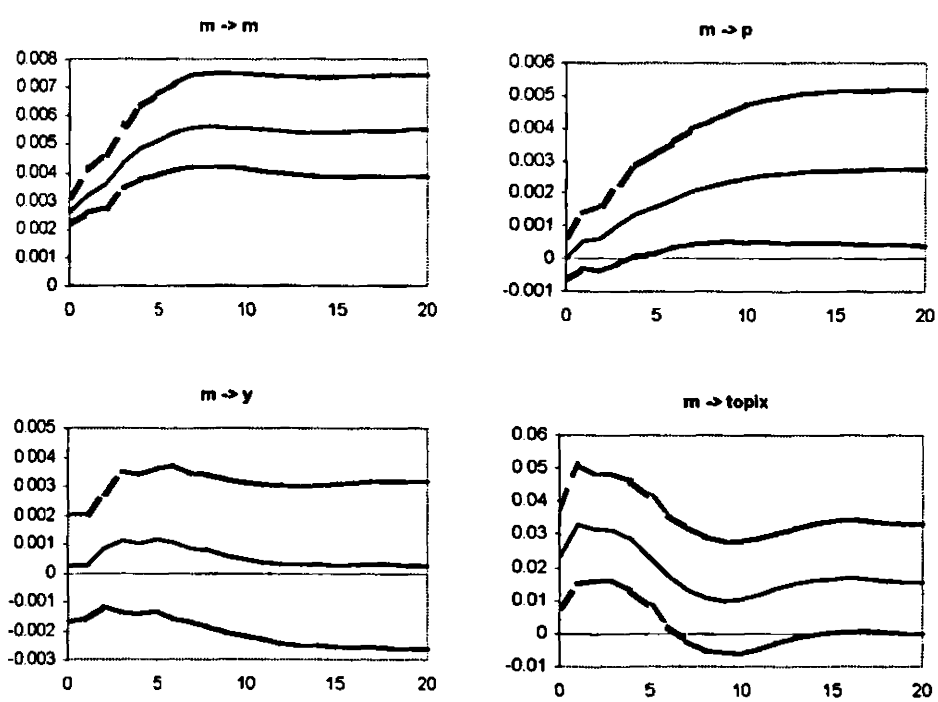


Figure 3. Impulse responses of nominal broadly defined liquidity (m), price level (p), real output (y) and real share prices ($topix$) to nominal liquidity shocks.

As illustrated by Figure 3, a shock to nominal money leads to a permanent and significant increase in the nominal money stock. Interestingly, it also leads to a permanent and significant increase in the price level. The point estimates of the impulse responses suggest a multiplier of 0.55 between nominal money and prices, even if, due to the broadness of the confidence intervals, a one-to-one impact cannot be ruled out. The impact of a nominal money shock on output is positive but statistically insignificant, whereas the effect on share prices is positive and significant in the short run. As a robustness test, we considered an alternative identification scheme where output shocks could have a contemporaneous effect on all the other variables, and prices were contemporaneously affected by shocks to output only. Nominal money was ordered third in the system, while share prices were assumed to be the most responsive variable to changes in economic conditions, as before. We found that the model dynamics remained robust to such an alternative identification scheme. Overall, the wide confidence bands for many of the shocks are likely to result from the relatively low estimation precision of our short sample. Estimating a real liquidity demand system may be then regarded as a possible method of increasing estimation precision, in addition to being theoretically more interesting.

Our findings can be compared with those of the Bank of Japan (2003), who evaluated the lead/lag relationships between nominal broadly defined liquidity, indicators for prices (CPI, land and share prices) and real output. This was accomplished by Granger-causality tests and by examining the cross-correlation coefficients between the growth rates of nominal money stock, prices and output. The BOJ found broadly defined liquidity to lead economic activity, both nominal and real GDP, but only in a sample ending in 1997Q3. Somewhat surprisingly, Granger-causality was reported to be running from asset prices to broad money, but no Granger-causal relationship was detected between broad money and consumer prices. It is clear that our results about the dynamics between the variables differ from those reported by the BOJ, while it should be kept in mind that both the estimation samples and the methodologies of the two studies are different. In particular, the Bank of Japan (2003) analyzed the lead/lag relationships, whereas we are interested in causality as determined in an impulse response analysis. Moreover, causality defined as in Granger (1969) concerns the predictive ability of a variable, not economic causality.

It may be of interest to conduct a brief analysis of cross-correlation coefficients between broad money and price indicators also in our paper, before proceeding to impulse response analysis in the context of the real liquidity system. Two main motivations are prominent. Firstly, one could argue in favour of using the growth rates of the real instead of the nominal money stock. This is especially the case if we adopt the view of Batini (2002) of money being a quantity-side measure of policy in the analysis. Arai and Hoshi (2004) similarly argued that an inference regarding policy stance is difficult using nominal growth rates of monetary aggregates - single digit growth may indicate tight policy when inflation rates are in double digits.[12] Secondly, it is easy to accommodate different measures of prices into this analysis, as was indeed done in the investigation by the Bank of Japan (2003). In particular, the only asset price indicator in our estimated VEC system is the share price index. We investigate the issue by calculating the cross-correlation coefficients for the variables under study, using a measure of real broad money growth lagged by k months, in order to evaluate its information content regarding future consumer and asset price developments. This kind of approach has been used to investigate the degree of inflation persistence, in the context of the time it takes for monetary settings to have the maximum effect on inflation. The approach was pioneered by Friedman (1972), and was subsequently used in studies by Friedman and Schwartz (1982), Batini and Nelson (2001) and Batini (2002). The results are listed in Table 4 below, with the maximum coefficients displayed together with the number of periods k when the maximum value is obtained. A maximum value of k=14 is assumed, corresponding to 3.5 years of quarterly data.

Price indicator	correlation	k	p-value
cpi	0.74	13	0.00
cpi (using M2+CDs)	0.68	13	0.00
topix(nominal)	0.55	0	0.00
residential land price	0.81	5	0.00
land price (all areas)	0.86	7	0.00

Table 4. Cross-correlation coefficients, annual growth rates of broadly defined liquidity and various price indicators.

Note: k indicates the number of lags of money growth whereby the reported maximum coefficient is obtained. p-values denote the probability of the null hypothesis that the coefficients are not significantly different from zero. Consumer prices are adjusted using the estimate by Sekine (2001) that a consumption tax hike would have caused a shift of 1.4% in the price index in 1997Q2 (consumer price index less fresh food).

[12] Interestingly, Towe (1998) found financial variables (including broad money M2+CDs) defined in real terms to perform much better in a forecasting exercise for economic activity than the same variables expressed in nominal terms.

While the maximum cross-correlation coefficient of broadly defined liquidity with the consumer price index obtains with over a 3 year lag, suggesting considerable inflation (deflation!) persistence in the Japanese economy, its magnitude and statistical significance are high. Then, a change in monetary conditions in the economy could indeed be captured by the growth rate of the broadly defined liquidity aggregate. The cross-correlation coefficient between consumer prices and broadly defined liquidity is slightly higher in absolute terms than when using M2+CDs as the money stock variable, and interestingly, the highest value is obtained at the same number of lags. As expected, the persistence of asset prices appears to be lower, with the maximum cross-correlation obtained contemporaneously for shares and with 5 lags for residential land prices. The coefficients are rather high and statistically significant, most notably so for overall land prices, again indicating the importance of broadly defined monetary conditions for asset price developments.[13] Moreover, it is not surprising that land prices react somewhat more slowly to an increase in the money stock than the prices of equities.

A final caveat is in order. The analysis in this subsection about the dynamics and causality between the money stock and various economic indicators has been conducted (as is conventional) using revised data for the investigation period. When data about the money stock is needed to provide information about economic conditions, only preliminary data may be available for the use of the central bank. In the case of broadly defined liquidity, the preliminary figures are reported of being of lower quality when compared to the available information on M2+CDs. According to the Bank of Japan (2004), the average absolute deviations in year-to-year percentage changes between the preliminary and final figures for fiscal 2002 were 0.05% for M2+CDs and 0.27% for broadly defined liquidity, respectively.

[13] The importance of land prices in the transmission of monetary policy in Japan was analysed by Ogawa (2000), who found that the fall in land value had weakened the effectiveness of policy considerably through rising external finance premiums in the 1990s. Kwon (1998) reached a similar conclusion about the important propagating role of Japanese land prices in the monetary transmission mechanism.

3.5.2 Identification with Long-Run Restrictions

From the investigation of relationships between the variables by simple correlation coefficients, we now return to examining the system dynamics by means of an impulse response analysis. When such a methodology was employed in the previous section, identification was achieved by contemporaneous restrictions only. An alternative identification scheme, discussed next, explicitly takes into account the cointegration properties of the system, relying predominantly on long-run restrictions to identify the model. Here, inference is obtained in the context of the real money demand system, being arguably more interesting from a theoretical perspective. Using the approach of the common trends literature (see e.g. King *et al.*, 1991), in our system with one cointegration relation, one shock will have transitory effects. Our stationary cointegration relationship was considered a plausible real money demand relation, justifying a restriction of a structural money demand shock to only have transitory effects on the other variables. Similarly, in the structural money demand system of Coenen and Vega (1999) for the euro area, money demand shocks only had transitory effects in the system. With our ordering of the variables, $(m-p)_t, y_t, topix_t$, the following long-run impact matrix ΞB is obtained:

$$\begin{bmatrix} 0 & * & * \\ 0 & * & * \\ 0 & * & * \end{bmatrix} \tag{3.17}$$

Again, asterisks denote the unrestricted elements. The second and third structural shocks are called supply and stock price shocks, respectively. Because the long-run impact matrix ΞB is of a reduced rank, only two linearly independent restrictions have been imposed. It could be appealing to further distinguish supply shocks from stock price shocks by assuming that the latter have no long-run impact on real output, resulting in a just-identified model. Since money demand shocks have a zero impact on real output as well, these restrictions would jointly amount to a Blanchard-Quah type of restriction of long-run demand neutrality. However, estimating the structural model with such an identification scheme yielded the counterintuitive result of a structural share price shock having a negative significant impact on output in the short run. This finding is identical to that of Miyao (2002) in an analysis of the identification of Japanese monetary policy shocks. As the author suggested, such a result could reflect an incorrectly imposed restriction in the case where the long-run effects between the variables are in fact large in magnitude. Indeed, theory suggests several reasons why output may increase as a result of higher share prices. Firstly, higher share prices increase the value of a firm relative to the replacement cost of its capital stock (higher Tobin's q; Tobin, 1969), and new investments can be financed with a relatively small issuance of new shares. Secondly, a permanent increase in asset prices induces an increase in permanent income *via* the wealth effect, leading to higher current and future consumption (Modigliani, 1971). Thirdly, higher asset prices may lower the probability of consumers' financial distress, increasing expenditure in durables and housing. Finally, a greater net worth position of firms, or a decline in the external finance premium of borrowers, leads to an increase in real investment through an increase in bank lending (the balance-sheet effect of Bernanke and Gertler, 1989). Consequently, instead of imposing such a long-term restriction in order to just-identify the model, we opted for a contemporaneous restriction on the B matrix, whereby shocks to share prices do not have a contemporaneous impact on real output due to stickiness in production, consumption and investment plans. These considerations result in the following matrix for B and in a just-identified model:

$$\begin{bmatrix} * & * & * \\ * & * & 0 \\ * & * & * \end{bmatrix} \tag{3.18}$$

It is possible to express the long-run restrictions as linear restrictions (see Breitung *et al.*, 2004) in the estimation procedure, while the general methodology follows the one utilized in the framework of contemporaneous restrictions. Our estimate for the contemporaneous impact matrix is found to be as follows, with the bootstrapped standard errors with 2,000 replications in parentheses:

$$\begin{bmatrix} \begin{matrix}0.0012\\(0.0003)\end{matrix} & \begin{matrix}0.0001\\(0.0003)\end{matrix} & \begin{matrix}0.0023\\(0.0003)\end{matrix} \\ \begin{matrix}-0.0001\\(0.0001)\end{matrix} & \begin{matrix}0.0075\\(0.0007)\end{matrix} & 0.0000 \\ \begin{matrix}-0.0419\\(0.0059)\end{matrix} & \begin{matrix}-0.0049\\(0.0085)\end{matrix} & \begin{matrix}0.0473\\(0.0082)\end{matrix} \end{bmatrix} \tag{3.19}$$

The impact response of the variable that may be considered to be most closely related to a particular structural shock is positive and significant in our system, as indicated by the diagonal elements of the matrix. The zero coefficient on the contemporaneous impact of share price shocks on real output is the outcome of our only contemporaneous restriction. Information beyond the contemporaneous dynamics between the system variables can be obtained by means of structural impulse response analysis, illustrated in Figure 4 below. Again, the Hall percentile 95% bootstrap confidence intervals were used.

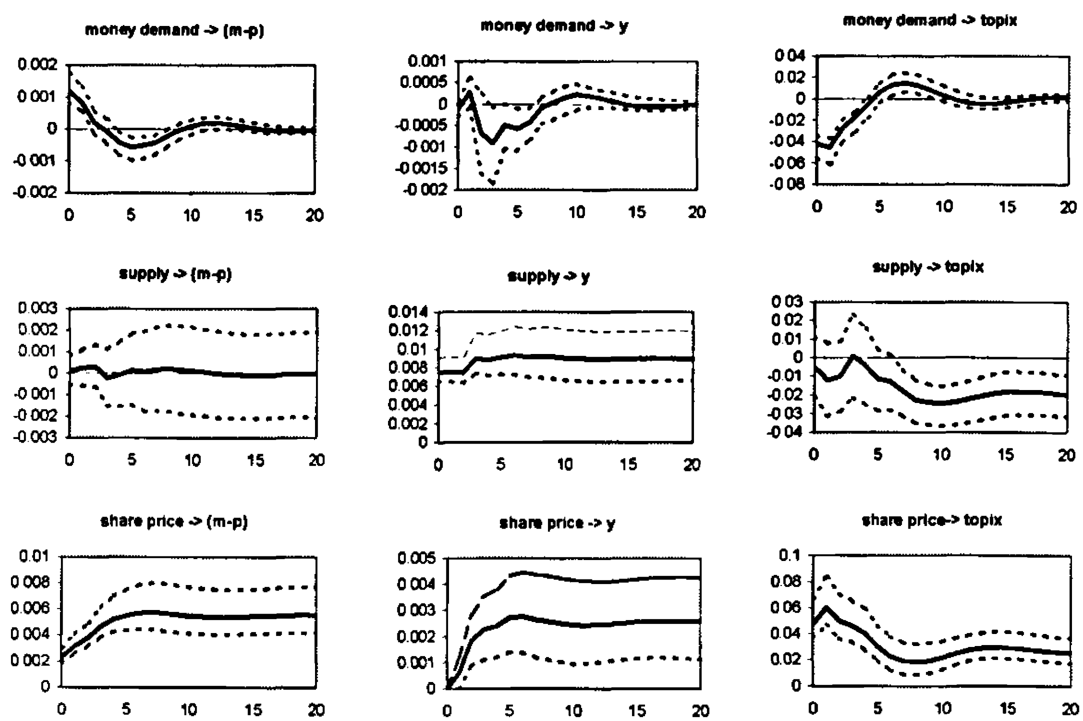


Figure 4. Structural impulse responses of real broadly defined liquidity $(m-p)$, real output (y) and real share prices $(topix)$ to money demand, supply and share price shocks.

A supply shock has an insignificant impact on the real liquidity stock. The negative effect of the supply shock on share prices may appear surprising, but if share prices are determined by other than real factors (an asset price bubble by definition serves as an example), such a finding is plausible. A structural share price shock leads to a significant and permanent increase in the money stock. Similarly, and as predicted by theory, the impact of a share price shock on real output is positive and significant in the long run. A counterintuitive impulse response is the one of real output to a structural money demand shock, as it is negative in the short run before finally turning positive (albeit insignificantly so) at 8 quarters. In the short run a money demand shock leads to a fall in the share price index, possibly due to portfolio adjustment from equities to broad money by the Japanese money holding sector.

The two identification schemes, employed in the context of models for real and nominal broadly defined liquidity, allow for a comparison of the resulting dynamics. The effects of share price shocks are similar for both identification schemes, even if the impact on output is more clearly statistically significant in the model for real liquidity demand. Also, the effects of supply shocks bear close resemblance to shocks to real output in the model with contemporaneous restrictions, with a statistically insignificant impact on both the real and nominal money stock. It is in the context of the money shock that the two identification schemes afford different outcomes. A money demand shock (expressed in terms of real broadly defined liquidity) has a transitory negative impact on output, whereas the effect of a shock to the nominal money stock on output was found to be positive but insignificant. Similarly, their impacts on share prices are different; in the case of a shock to real liquidity demand, the negative coefficient was suggested to reflect portfolio adjustment away from equities by the money-holding sector.

This section examined the model dynamics in the context of an impulse response analysis, using a structural vector error correction model with both contemporaneous and long-run restrictions. Interestingly in the context of the persistent Japanese deflation problem, causal dynamics were detected between the nominal money stock and both the consumer and share price indices. The usefulness of the broad money stock as an information variable was confirmed in the investigation of cross-correlation coefficients between broadly defined liquidity and various price indicators.

3.6 Conclusions

This paper set out to examine the demand for broadly defined liquidity in Japan during 1981-2004. Our analysis established a relatively stable and an economically meaningful demand relationship for this rarely-investigated monetary aggregate in a cointegration framework. This is a significant finding, as the estimation sample was characterized by anxieties in the financial system and the start of the deflationary environment in the Japanese economy. Our results are in contrast to some of the earlier literature which suggested that the relationship between real economy and the broad money stock has become unstable in recent years, or even that an economically meaningful relationship could not be established for broadly defined liquidity in a cointegration framework. We confirmed the significance of share prices for Japanese broad money demand, with the wealth effect of equity values prevailing over the substitution effect. Little evidence of important currency substitution phenomena was detected. Finally, the investigation of system dynamics was based on impulse response analysis in the context of a structural vector error correction (SVEC) model for the Japanese broadly defined liquidity.

We do not suggest that broadly defined liquidity should be targeted by the BOJ, or even that such a strategy would be feasible for a monetary authority. Rather, broadly defined liquidity seems to satisfy the requirements of an important information variable, as a relatively stable long-run relationship with economic activity was detected. Moreover, a causal relationship between broad liquidity and consumer prices could not be rejected in the nominal liquidity demand framework. A similar finding is the high cross-correlation coefficient between (lagged) broad money growth and movements in consumer prices. This is of particular interest, given the significant persistence of deflation in the Japanese economy. The recent acceleration in economic growth which commenced in late 2003, while alleviating negative price pressures, has so far failed to pull the economy fully out of the deflationary trap. Even with possible instabilities in the money multiplier, and the fact that high monetary base growth has not been reflected in broad money growth in an equivalent magnitude, the information value of broad money is not invalidated. Similarly, the causality of broadly defined liquidity for share prices is in line with recently revived interest in the information value of broad money growth for asset price developments. The detected importance of equity values for broad money demand in Japan seems rather logical considering the asset price bubble in this economy that led to stubbornly weak economic growth - "the great recession" of the 1990s.

Admittedly, not all aspects of the empirical analysis are optimal. The stability of the system during the bursting of the asset price bubble may not be optimally guaranteed by our shift dummy variable for the liquidity variable in first differences. Moreover, the negative short-run impact of a structural money demand shock on real output in the context of the SVEC analysis is counterintuitive. However, our wish is to revive interest in the information value of broad monetary aggregates for a central bank operating during a deflationary era.

Bibliography

[1] Amano, R. A., Wirjanto, T. S., 2000. On the Stability of Long-Run M2 Demand in Japan. The Japanese Economic Review 51 (4), 536-543.

[2] Amisano, G., Giannini, C., 1997. Topics in Structural VAR Econometrics. Springer-Verlag, Berlin.

[3] Arai, Y., Hoshi, T., 2004. Monetary Policy in the Great Recession. In: Hutchison, M., Westermann, F. (Eds.), Economic Stagnation in Japan, CesIFO MIT Conference Volume.

[4] Artis, M., Beyer, A., 2004. Issues in Money Demand: The Case of Europe. Journal of Common Market Studies 42 (4), 717-736.

[5] Bahmani-Oskooee, M., 2001. How Stable is M2 Money Demand Function in Japan? Japan and the World Economy 13 (4), 455-461.

[6] Bank of Japan, 2003. The Role of the Money Stock in Conducting Monetary Policy. Bank of Japan Quarterly Bulletin May 2003, 1-52.

[7] Bank of Japan, 2004. Guide to Japan's Money Stock Statistics. Publication of the Research and Statistics Department, Bank of Japan.

[8] Batini, N., 2002. Euro Area Inflation Persistence. ECB Working Paper no. 201.

[9] Batini, N., Nelson, E., 2001. The Lag from Monetary Policy Actions to Inflation: Friedman Revisited. International Finance 4 (3), 381-400.

[10] Bayoumi, T., 2001. The Morning After: Explaining the Slowdown in Japanese Growth in the 1990s. Journal of International Economics 53, 241-259.

[11] Benkwitz, A., Lütkepohl, H., Wolters, J., 2001. Comparison of Bootstrap Confidence Intervals for Impulse Responses of German Monetary Systems. Macroeconomic Dynamics 5, 81-100.

[12] Bernanke, B. S., Gertler, M., 1989. Agency Costs, Net Worth and Business Fluctuations. American Economic Review 79, 14-31.

[13] Breitung, J., Brüggemann, R., Lütkepohl, H., 2004. Structural Vector Autoregressive Modeling and Impulse Responses. In: Lütkepohl, H., Krätzig, M. (Eds.), Applied Time Series Econometrics. Cambridge University Press, Cambridge.

[14] Brüggemann, R., Lütkepohl, H., 2004. Practical Problems with Reduced Rank ML Estimators for Cointegration Parameters and a Simple Alternative. European University Institute Working Paper no. 2004/20.

[15] Candelon, B., Lütkepohl, H., 2001. On the Reliability of Chow-Type Tests for Parameter Constancy in Multivariate Dynamic Models. Economics Letters 73, 155-160.

[16] Coenen, G., Vega, J. L., 1999. The Demand for M3 in the Euro Area. ECB Working Paper no. 6.

[17] Corsetti, G., Pesenti, P., 1999. Stability, Asymmetry, and Discontinuity: The Launch of European Monetary Union. Brookings Papers on Economic Activity 1999 (2), 295-372.

[18] ECB, 2001. Monthly Bulletin, December 2001. European Central Bank Publications.

[19] Ericsson, N. R., 1999. Empirical Modeling of Money Demand. In: Lütkepohl, H., Wolters, J. (Eds.), Money Demand in Europe. Physica-Verlag, Heidelberg.

[20] Friedman, M., 1972. Have Monetary Policies Failed? American Economic Review (Papers and Proceedings) 62, 11-18.

[21] Friedman, M., 1988. Money and the Stock Market. Journal of Political Economy 96, 221-245.

[22] Friedman, M., Schwartz, A. J., 1982. Monetary Trends in the United States and the United Kingdom: Their Relation to Income, Prices, and Interest Rates, 1867-1975. University of Chicago Press for NBER.

[23] Gerlach, S., 2004. The Two Pillars of the European Central Bank. Economic Policy 40, 389-439.

[24] Goodfriend, M., 2000. Overcoming the Zero Bound on Interest Rate Policy. Journal of Money, Credit and Banking 32 (4), 1007-1035.

[25] Granger, C. W. J., 1969. Investigating Causal Relations by Econometric Models and Cross-Spectral Methods. Econometrica 37, 424-438.

[26] Hansen, H., Johansen, S., 1999. Some Tests for Parameter Constancy in Cointegrated VAR-Models. Econometrics Journal 2, 306-333.

[27] Hetzel, R. L., 2003. Japanese Monetary Policy and Deflation. Federal Reserve Bank of Richmond Economic Quarterly 89 (3), 21-52.

[28] Ho, C., Ma., G., McCauley, R. N., 2005. Trading Asian Currencies. BIS Quarterly Review, March 2005, 49-58.

[29] Issing, O., 2002. Why Stable Prices and Stable Markets Are Important and How They Fit Together. Speech at the First Conference of the Monetary Stability Foundation, 5 December 2002.

[30] Ito, T., Mishkin, F. S., 2004. Two Decades of Japanese Monetary Policy and the Deflation Problem. NBER Working Paper no. 10878.

[31] Iwaisako, T., 2003. Household Portfolios in Japan. NBER Working Paper no. 9647.

[32] Jang, K., Ogaki, M., 2003. The Effects of Japanese Monetary Policy Shocks on Exchange Rates: A Structural Vector Error Correction Model Approach. Bank of Japan Monetary and Economic Studies, February 2003.

[33] JMulTi, 2004. Analysing Multiple Time Series with JAVA. Chair of Econometrics, Humboldt University Berlin. Available at www.jmulti.com.

[34] Johansen, S., 1995. Likelihood Based Inference in Cointegrated Vector Autoregressive Models. Oxford University Press, Oxford.

[35] Kameda, S., Kyoso, K., Yoshida, T., 1998. Searching for the Long-Run Relations between Monetary Aggregates and Other Macroeconomic Variables in Japan: A Vector Error Correction Approach. Bank of Japan Working Paper no. 98-2.

[36] Kimura, T., 2001. The Impact of Financial Anxieties on Money Demand in Japan. In: Klöckers, H-J., Willeke, C. (Eds.), Monetary Analysis: Tools and Applications. European Central Bank Publications, Frankfurt am Main.

[37] Kimura, T., Kobayashi, H., Muranaga, J., Ugai, H., 2002. The Effect of the Increase in Monetary Base on Japan's Economy at Zero Interest Rates: An Empirical Analysis. IMES Discussion Paper no. 2002-E-22.

[38] King, R. G., Plosser, C. I., Stock, J. H., Watson, M., 1991. Stochastic Trends and Economic Fluctuations. American Economic Review 81 (4), 819-840.

[39] Kwon, E., 1998. Monetary Policy, Land Prices, and Collateral Effects on Economic Fluctuations: Evidence from Japan. Journal of the Japanese and International Economies 12, 175-203.

[40] Lanne, M., Lütkepohl, H., Saikkonen, P., 2002. Comparison of Unit Root Tests for Time Series with Level Shifts. Journal of Time Series Analysis 23 (6), 667-685.

[41] Lütkepohl, H., 2004. Vector Autoregressive and Vector Error Correction Models. In: Lütkepohl, H., Krätzig, M. (Eds.), Applied Time Series Econometrics. Cambridge University Press, Cambridge.

[42] Lütkepohl, H., Wolters, J., 1999. A Money Demand System for German M3. In: Lütkepohl, H., Wolters, J. (Eds.), Money Demand in Europe. Physica-Verlag, Heidelberg.

[43] Masuch, K., Nicoletti-Altimari, S., Pill, H., Rostagno, M., 2003. The Role of Money in Monetary Policy Making. In: Background Studies for the ECB's Evaluation of its Monetary Policy Strategy. ECB Publications.

[44] McCallum, B. T., 1988. Robustness Properties of a Rule for Monetary Policy. Carnegie-Rochester Conference Series on Public Policy 29 (Autumn), 173-203.

[45] McCallum, B. T., 2003. Japanese Monetary Policy, 1991-2001. Federal Reserve Bank of Richmond Economic Quarterly 89 (1), Winter 2003, 1-31.

[46] Meltzer, A., 1999. The Transmission Process. Paper Presented at the Bundesbank Conference on the Transmission Process: Recent Developments and Lessons for Europe.

[47] Miyao, R., 2002. The Effects of Monetary Policy in Japan. Journal of Money, Credit, and Banking 34 (2), 376-392.

[48] Miyao, R., 2005. Use of the Money Supply in the Conduct of Japan's Monetary Policy: Re-Examining the Time-Series Evidence. The Japanese Economic Review 56 (2), 165-187.

[49] Modigliani, F., 1971. Monetary Policy and Consumption. In: Consumer Spending and Monetary Policy: Linkages. Federal Reserve Bank of Boston, Boston.

[50] Morsink, J., Bayoumi, T., 2000. Monetary Policy Transmission in Japan. In: Bayoumi, T., Collyns, C. (Eds.), Post-Bubble Blues: How Japan Responded to Asset Price Collapse. International Monetary Fund, Washington DC.

[51] Nelson, E., 2003. The Future of Monetary Aggregates in Monetary Policy Analysis. Journal of Monetary Economics 50, 1029-1059.

[52] Ogawa, K., 2000. Monetary Policy, Credit, and Real Activity: Evidence from the Balance Sheet of Japanese Firms. Journal of the Japanese and International Economies 14, 385-407.

[53] Pantula, S. G., 1989. Testing for Unit Roots in Time Series Data. Econometric Theory 5, 256-271.

[54] Ploberger, W., Krämer, W., Kontrus, K., 1989. A New Test for Structural Stability in the Linear Regression Model. Journal of Econometrics 40, 307-318.

[55] Saikkonen, P., Lütkepohl, H., 2000. Testing for the Cointegrating Rank of a VAR Process with Structural Shifts. Journal of Business & Economic Statistics 18, 451-464.

[56] Sekine, T., 1998. Financial Liberalization, the Wealth Effect, and the Demand for Broad Money in Japan. Bank of Japan Monetary and Economic Studies, May 1998, 35-56.

[57] Sekine, T., 2001. Modeling and Forecasting Inflation in Japan. IMF Working Paper no. 01/82.

[58] Tavlas, G. S., 1996. Currency Substitution and the International Demand for Yen. In: Mizen, P., Pentecost, E. J. (Eds.), The Macroeconomics of International Currencies - Theory, Policy and Evidence. Edward Elgar, Cheltenham.

[59] Tobin, J., 1969. A General Equilibrium Approach to Monetary Theory. Journal of Money, Credit, and Banking 1, 15-29.

[60] Towe, C., 1998. Structural Change and Information Content of Financial Variables for Monetary Policy. In: Aghevli, B. B., Bayoumi, T., Meredith, G. (Eds.), Structural Change in Japan: Macroeconomic Impact and Policy Challenges. International Monetary Fund, Washington DC.

[61] Tsukuda, Y., Miyakoshi, T., 1998. Granger Causality between Money and Income for the Japanese Economy in the Presence of a Structural Change. The Japanese Economic Review 49 (2), 191-209.

[62] Warne, A., 2005. Structural VAR - version 0.32. Available at http://www.texlips.net/svar/.

[63] Yoshida, T., 1990. On the Stability of the Japanese Money Demand Function: Estimation Results Using the Error Correction Model. Bank of Japan Monetary and Economic Studies 8 (1), 1-48.

[64] Yoshida, T., Rasche, R. H., 1990. The M2 Demand in Japan: Shifted and Unstable? Bank of Japan Monetary and Economic Studies 8 (2), 9-30.

3.7 Appendices

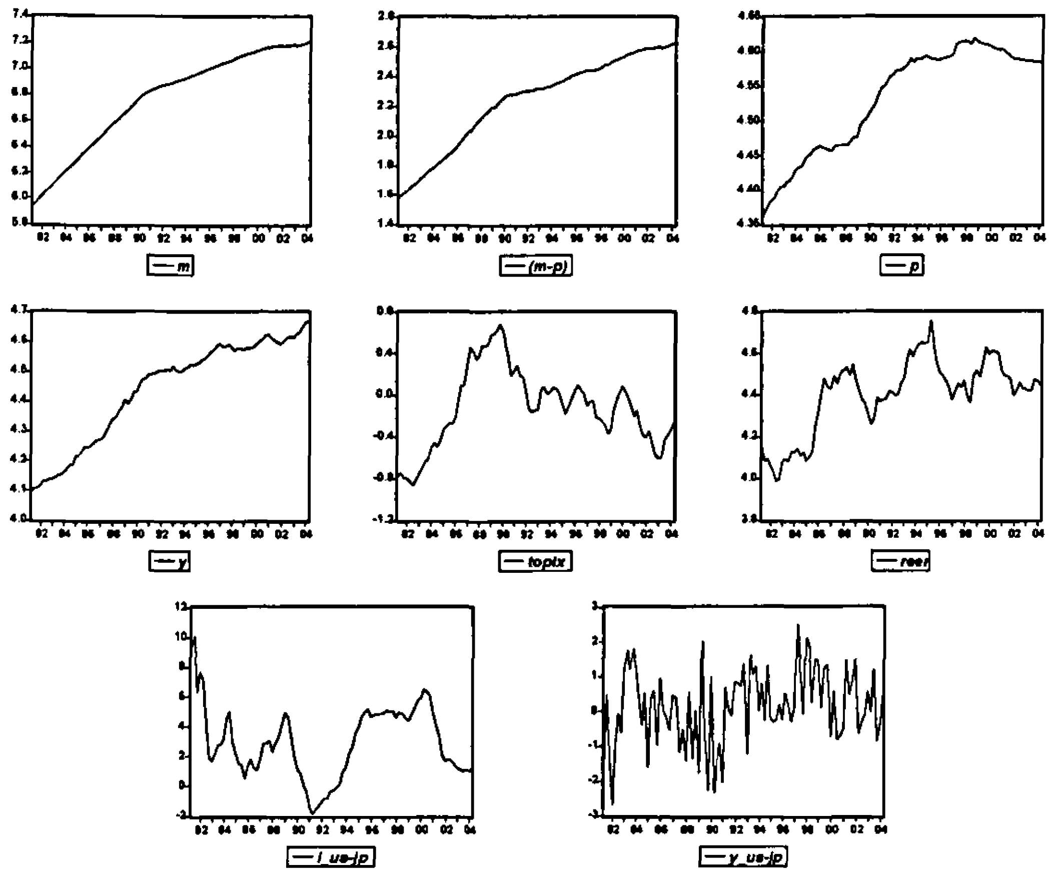
Appendix A

Data Sources

The following variables were obtained from the OECD Main Economic Indicators Database, vol 2004, release 10: Gross domestic product, volume (both for the US and Japan), broadly defined liquidity, the consumer price index, the share price index TOPIX, the real effective exchange rate, the 3-month CD rate (both for the US and Japan).

All the series are at a quarterly frequency. All series except the real exchange rate and the interest rates are seasonally adjusted, as reported by the OECD. The consumer price index was seasonally adjusted by the author using a Census X-11 seasonal adjustment procedure. All variables except the GDP growth and interest rate differentials between the US and Japan were transformed into logarithms.

Figures of Series Used in Estimation



Appendix B: Unit Root and Cointegration Tests

Augmented Dickey-Fuller Test

Series	Det. term	Lagged differences	Test stat.
$\Delta^2 y$	c	5 (AIC,HQ,SC)	-2.96**
Δy	c	2 (AIC,HQ,SC)	-3.28**
y	c,t	3 (AIC)	-1.57
y	c,t	0 (HQ,SC)	-0.83
$\Delta^2 p$	c	5 (AIC, HQ)	-2.08
$\Delta^2 p$	c	4 (SC)	-1.73
Δp	c	2 (AIC,HQ,SC)	-2.88**
p	c,t	3 (AIC,HQ)	-0.46
p	c,t	0 (SC)	0.36
$\Delta^2 topix$	c	6 (AIC,HQ)	-2.38
$\Delta^2 topix$	c	4 (SC)	-2.91**
$\Delta topix$	c	0 (AIC,HQ,SC)	-6.36***
$topix$	c,t	1 (AIC,HQ,SC)	-1.90
$\Delta^2 reer$	c	6 (AIC)	-3.15**
$\Delta^2 reer$	c	5 (HQ)	-4.33***
$\Delta^2 reer$	c	3 (SC)	-4.36***
$\Delta reer$	c	0 (AIC,HQ,SC)	-8.03***
$reer$	c,t	1 (AIC)	-1.94
$reer$	c,t	0 (HQ,SC)	-1.53
$\Delta^2 y^{US-JP}$	c	7 (AIC)	-5.61***
$\Delta^2 y^{US-JP}$	c	3 (HQ,SC)	-6.42***
Δy^{US-JP}	c	1 (AIC,HQ,SC)	-12.96***
y^{US-JP}	c	2(AIC,HQ,SC)	-3.23**
$\Delta^2 i^{US-JP}$	c	10 (AIC)	-1.77
$\Delta^2 i^{US-JP}$	c	3 (HQ,SC)	-4.92***
Δi^{US-JP}	c	0(AIC,HQ,SC)	-7.69***
i^{US-JP}	c	5 (AIC)	-1.88
i^{US-JP}	c	1 (HQ,SC)	-3.43**

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

The order specification criteria in parentheses: AIC=Akaike, HQ=Hannan-Quinn, SC=Schwarz-criteria.

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

All series in logarithms except y^{US-JP} and i^{US-JP}.

Maximum lag order set at 10, sample 1981Q2-2004Q2.

Unit Root Tests with Structural Break (Lanne *et al.*, 2002)

Series	Det. term	Lagged differences	Test stat.
$\Delta^2(m-p)$	$c, S90Q4$	9 (AIC)	-1.94
$\Delta^2(m-p)$	$c, S90Q4$	7 (HQ)	-2.40
$\Delta^2(m-p)$	$c, S90Q4$	0 (SC)	-3.36**
$\Delta(m-p)$	$c, S90Q3$	0 (AIC,HQ,SC)	-8.04***
$\Delta^2 m$	$c, S90Q3$	8 (AIC)	-1.57
$\Delta^2 m$	$c, S90Q3$	4 (HQ,SC)	-1.63
Δm	$c, S90Q3$	1 (AIC,HQ)	-3.20**
Δm	$c, S90Q3$	0 (SC)	-3.98***

* indicates significance at 10% level, ** at 5% and *** at 1% level.
The order specification criteria in parentheses: AIC=Akaike, HQ=Hannan-Quinn, SC=Schwarz-criteria.
c and t denote constant and trend as deterministic terms, respectively.
All series in logarithms.
Prefix S denotes date of shift dummy.
Maximum lag order set at 10, sample 1981Q2-2004Q2.

Saikkonen-Lütkepohl Cointegration Test

Series	Det. term	no. of lags	Coint. rank	test stat
$m, p, y, topix$	$c, S90Q3, S97Q2$	1 (AIC,HQ,SC)	0	157.58***
			1	20.18
			2	5.54
			3	0.72
$m, p, y, topix$	$c, t, S90Q3, S97Q2$	1 (AIC,HQ,SC)	0	97.32***
			1	37.60***
			2	6.11
			3	0.02
$(m-p), y, topix$	$c, S90Q3, S97Q2$	1 (AIC,HQ,SC)	0	128.76***
			1	6.87
			2	1.04
$(m-p), y, topix$	$c, t, S90Q3, S97Q2$	10 (AIC)	0	32.74**
			1	4.52
			2	0.09
$(m-p), y, topix$	$c, t, S90Q3, S97Q2$	1 (HQ,SC)	0	73.66***
			1	2.61
			2	0.05
$(m-p), y, topix$ [82Q2 - 04Q2]	$c, t, S90Q3, S97Q2$	2 (AIC)	0	54.44***
			1	2.06
			2	0.13
$(m-p), y, topix$ [82Q2 - 04Q2]	$c, t, S90Q3, S97Q2$	1 (HQ,SC)	0	93.29***
			1	2.39
			2	0.22
$(m-p), y, topix$ [81Q2 - 03Q2]	$c, t, S90Q3, S97Q2$	2 (AIC)	0	33.53***
			1	5.81
			2	1.13
$(m-p), y, topix$ [81Q2 - 03Q2]	$c, t, S90Q3, S97Q2$	1 (HQ,SC)	0	61.35***
			1	2.20
			2	0.18
$(m-p), y$	$c, t, S90Q3, S97Q2$	1 (AIC,HQ,SC)	0	21.70***
			1	0.26

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

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

The order specification criteria in parentheses: AIC=Akaike, HQ=Hannan-Quinn, SC=Schwarz-criteria.

Prefix S denotes date of shift dummy.

Alternative estimation sample denoted in brackets.

Saikkonen-Lütkepohl Cointegration Test (continued)

Series	Det. term	no. of lags	Coint. rank	test stat
$(m-p), topix$	$c, t, S90Q3, S97Q2$	10 (AIC)	0	11.69
			1	0.32
$(m-p), topix$	$c, t, S90Q3, S97Q2$	2 (HQ)	0	18.28**
			1	0.69
$(m-p), topix$	$c, t, S90Q3, S97Q2$	1 (SC)	0	34.30***
			1	0.27
$y, topix$	$c, t, S97Q2$	2 (AIC,HQ)	0	26.01***
			1	1.86
$y, topix$	$c, t, S97Q2$	1 (SC)	0	32.06***
			1	3.33

* indicates significance at 10%, ** at 5% and *** at 1% level.
c and t denote constant and trend as deterministic terms, respectively.
The order specification criteria in parentheses: AIC=Akaike, HQ=Hannan-Quinn, SC=Schwarz-criteria.
Prefix S denotes date of shift dummy.

Appendix C
Misspecification Tests

Real liquidity demand system

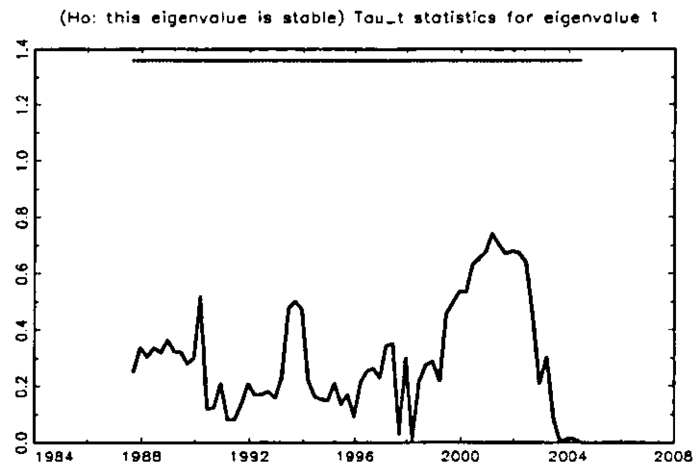
Q^*_{16}	106.26 [0.94]
LM_5, LM_4, LM_1	48.11 [0.35], 44.75 [0.15], 2.48 [0.98]
JB (eqs. 1, 2, 3)	25.60 [0.00] 2.93 [0.23] 0.30 [0.86]
$ARCH_{LM}(16)$(eqs. 1, 2, 3)	26.73 [0.04] 15.59 [0.48] 10.60 [0.83]

Nominal liquidity demand system

Q^*_{16}	203.92 [0.92]
LM_5, LM_4, LM_1	75.85 [0.61], 64.84 [0.45], 4.86 [1.00]
JB (eqs. 1, 2, 3, 4)	15.01 [0.00] 4.33 [0.11] 3.30 [0.19] 0.02 [0.99]
$ARCH_{LM}(16)$(eqs. 1, 2, 3, 4)	25.15 [0.07] 10.41 [0.84] 13.25 [0.65] 10.98 [0.81]

Note: p-values in brackets.
Q* denotes the adjusted Portmanteau test statistic for autocorrelation (conducted only for models without exogenous variables).
LM is the Lagrange multiplier type test statistic for autocorrelation.
JB is the Jarque-Bera test for nonnormality.
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.

Appendix D
Stability Tests



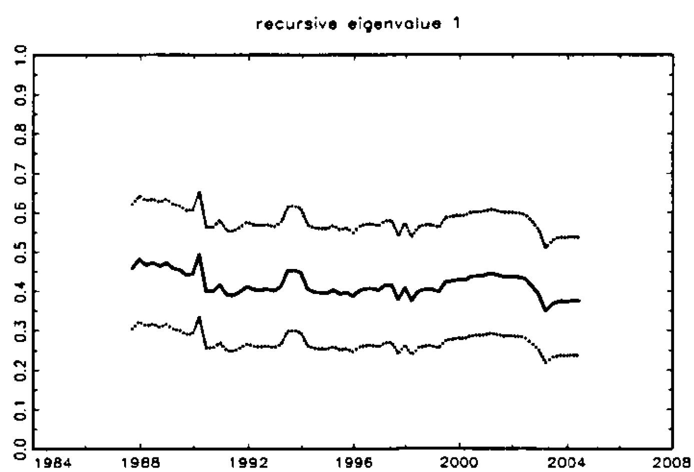


Figure: Recursive eigenvalue analysis of the VEC model for real broadly defined liquidity.

break date	chow_bp	boot p-val	chow_ss	boot p-val	break date	chow_bp	boot p-val	chow_ss	boot p-val
1988 Q3	47.59	0.243	17.25	0.823	1993 Q4	48.27	0.169	16.79	0.909
1988 Q4	50.86	0.177	17.48	0.841	1994 Q1	49.49	0.183	17.08	0.898
1989 Q1	52.55	0.143	17.88	0.837	1994 Q2	41.86	0.357	13.06	0.97
1989 Q2	54.58	0.103	18.47	0.834	1994 Q3	40.06	0.415	11.55	0.99
1989 Q3	47.36	0.185	13.67	0.926	1994 Q4	41.61	0.386	12.55	0.979
1989 Q4	50.99	0.096	15.13	0.865	1995 Q1	43.34	0.311	14.17	0.931
1990 Q1	49.10	0.131	16.02	0.86	1995 Q2	42.08	0.373	14.80	0.949
1990 Q2	51.89	0.084	19.54	0.705	1995 Q3	42.42	0.356	15.54	0.907
1990 Q3	48.32	0.134	18.98	0.788	1995 Q4	42.95	0.361	15.58	0.903
1990 Q4	50.13	0.159	19.10	0.807	1996 Q1	43.84	0.296	16.23	0.857
1991 Q1	52.58	0.132	20.46	0.788	1996 Q2	43.86	0.316	15.60	0.885
1991 Q2	46.77	0.22	19.69	0.802	1996 Q3	45.80	0.264	17.27	0.822
1991 Q3	49.25	0.151	21.02	0.74	1996 Q4	43.18	0.355	17.59	0.832
1991 Q4	50.51	0.126	20.68	0.765	1997 Q1	44.28	0.342	18.06	0.801
1992 Q1	52.37	0.107	21.35	0.729	1997 Q2	44.13	0.366	21.94	0.64
1992 Q2	48.70	0.148	19.68	0.818	1997 Q3	41.29	0.454	19.18	0.723
1992 Q3	50.86	0.129	19.13	0.83	1997 Q4	41.03	0.449	17.77	0.781
1992 Q4	49.30	0.147	17.98	0.887	1998 Q1	40.03	0.505	16.63	0.764
1993 Q1	48.29	0.185	17.54	0.897	1998 Q2	41.89	0.467	19.98	0.58
1993 Q2	47.55	0.195	16.58	0.926	1998 Q3	49.56	0.24	12.17	0.743
1993 Q3	50.75	0.115	18.18	0.862					

Chow break point (chow_bp) and sample split (chow_ss) test statistics for benchmark model.
Bootstrapped p-values based on 1,000 replications

equation	test statistic	p-value
$(m-p)_t$	1.6350	0.1318
y_t	1.7098	0.1098
$topix_t$	1.9298	0.0694

Table: Recursive fluctuation tests.
p-values based on 4,999 bootstrapping replications.

Appendix E
Impulse Response Analysis

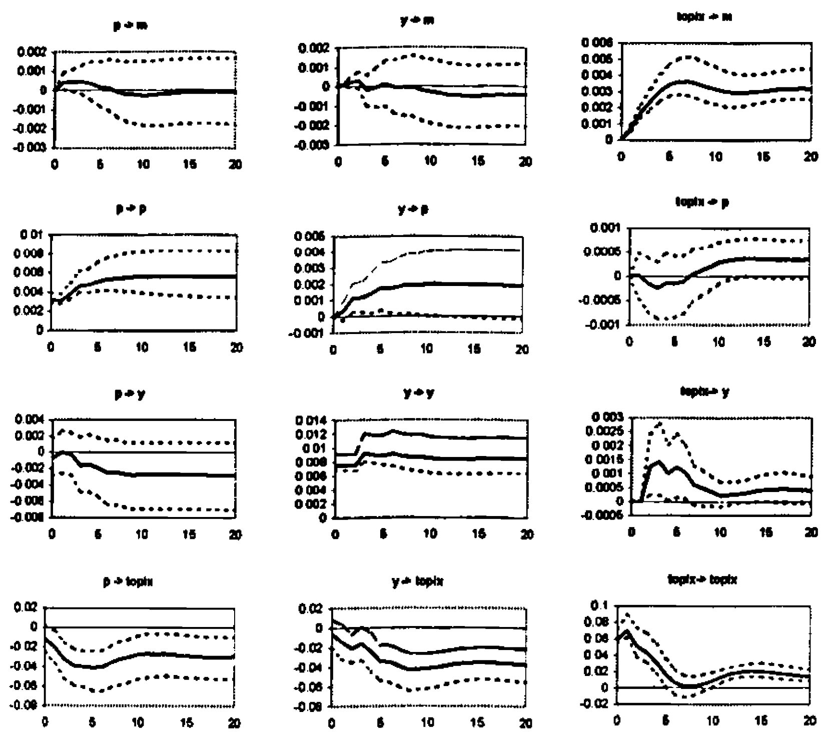


Figure: Impulse responses of nominal broadly defined liquidity (m), price level (p), real output (y) and real share prices ($topix$) to price shocks (first column), output shocks (second column) and share price shocks (third column).

Chapter 4

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

4.1 Introduction

The increasing occurrence of deflation episodes in both emerging and industrial economies has recently engaged the concern of policymakers, calling for an assessment of the possible causes and cures to falling prices. It is Japan's stagnation and deflation which has undoubtedly made most of the headlines and called for policy proposals from numerous economists. Yet the rates of deflation have been even higher in Hong Kong SAR where independent monetary policy is constrained by the currency board arrangement, while China's deflation has been experienced in an environment of a US dollar peg, capital controls and rapid economic growth. These strikingly different monetary regimes and economic environments make it interesting to evaluate the potency of the interest rate and exchange rate instruments for an economy operating during a deflationary era, with the ultimate aim of returning to price stability or an increasing price level path.

The aim of our paper is to examine the role of exchange and interest rate channels during the recent disinflation and deflation episodes in Asia. For this purpose, we estimate 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 frame of our study, we use nominal effective exchange rates in our analysis. While there have been previous studies of the causes of inflation and deflation in these economies, our approach has some novel features. By using a similar estimation framework for all three economies for the same time period, we are able to capture similarities and differences between the price adjustment processes for these economies. Further, to our knowledge no previous estimations of an open-economy SVAR model for a period of falling prices exist, allowing for the assessment of the relative importance of the exchange rate and interest rate shocks.

The evaluation of the exchange rate channel is rendered no less interesting by the fact that several suggestions to escape from a liquidity trap are centred around the exchange rate tool, perhaps most prominently in the "foolproof way" proposed by Svensson (2001, 2003). A historical example about the use of the exchange rate in order to end a period of deflation is President Roosevelt's policy of dollar depreciation in the 1930s (see Friedman and Schwartz, 1963). In contrast, the attraction of the interest rate instrument has been predominantly seen as lost once the zero bound is hit. In particular, the influential Taylor-type of rule (see Taylor, 1993) that characterizes the interest rate setting behaviour of a central bank as a linear function of inflation and the output gap, does not take into account the fact that interest rates are bounded by the zero floor.

We find that 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 the latter economy. Our results provide evidence about the role of external influences in the deflation episodes of these two economies. Moreover, while domestic interest rate policy is restricted by the currency board arrangement in Hong Kong, the high level of interest rates during the deflationary era may have represented a disadvantage of the dollar peg. Indeed, in both Hong Kong and Japan, the importance of the interest rate channel for price developments is found to be non-negligible. The results for Japan could be regarded as weakly supporting a strategy to depreciate the yen in order to escape from the liquidity trap, even if the required exchange rate movements are substantial. In China, where interest rates have not been an important monetary policy tool, neither exchange nor interest rate shocks significantly influence price developments.

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

4.2 Developments in the Three Economies

In this section, we provide a descriptive analysis of the development of consumer prices, interest and exchange rates in Japan, Hong Kong and China, mentioning some of the relevant literature concerning deflation in these economies.

In Japan, the economy saw consumer price inflation falling below zero in late 1995. Positive inflation re-emerged together with meagre growth in 1997, but after that deflation took hold again. All in all, the Japanese consumer price index has remained relatively stable during the time period of our study. This is evident in Figure 1 below, where we depict the annual consumer price inflation rates for all of the three economies.[1] The deflationary episode in Japan has been broad-based in nature, as the prices of major items in the consumer price index (food, clothing, transportation, and durable goods) have all registered declines (IMF, 2003). The nominal exchange rate strongly depreciated after its peak in 1995, having appreciated by 20 percent against the dollar between January and April, but fully reversing the climb already by September of the same year. In mid-1998, it rebounded and another appreciation peak emerged in 2000, as depicted in Figure 2 further below. Finally, the limitations for conventional monetary policy of relying on the interest rate channel have been well documented in the literature.[2] The zero interest rate policy was initiated by the Bank of Japan in early 1999; it was then briefly abandoned in August 2000 and re-introduced in March 2001. Other short-term rates beside the policy rate have also been hovering close to the zero bound (the behaviour of the 3-month CD rate used in our study is depicted in Figure 3 in the end of this section).

[1] The series for China only begins in 1994 due to reasons of data availability.

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

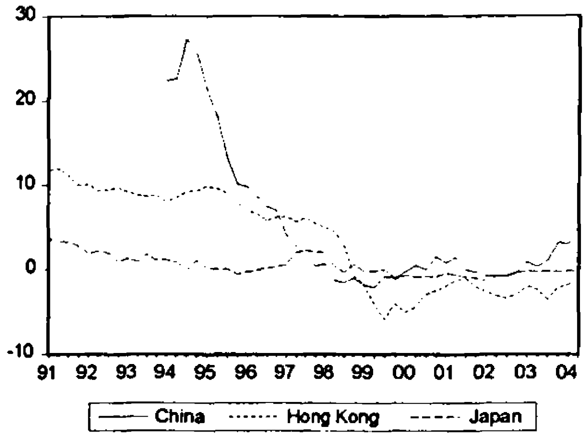


Figure 1. Consumer price inflation rates (year-on-year), China, Hong Kong and Japan, 1991-2004.

In Hong Kong, after experiencing a CPI inflation rate of 10% still in early 1995, the drop in the inflation rate was extremely rapid and deflation first emerged in late 1998. By mid-1999, deflation was already running at a yearly rate of 6%. Since then, the rate of deflation has moderated and positive annual inflation rates were finally recorded in the second half of 2004. Major contributors to deflation have been the cost of housing, food, durable goods and clothing. Schellekens (2003) showed that these items together contributed over 85% to overall deflation. In an empirical study, the author found that deflation in Hong Kong SAR was best explained by cyclical shocks, amplified by balance-sheet and wealth effects. In contrast, the role of price equalization with Mainland China was found to be small. Investigating inflation dynamics in Hong Kong, Genberg and Pauwels (2003) argued that import prices, wages and property prices were important sources of changes in marginal costs in Hong Kong. The nominal effective exchange rate of the Hong Kong dollar was steadily appreciating until 2002, notwithstanding a fall in 1998 at the time of the Asian financial crisis. After that, the fall in the US dollar has brought about a depreciation in the exchange rate. This can be seen in Figure 2 below. Meanwhile, independent monetary policy through the interest rate channel has been limited due to the currency board arrangement with the US dollar, with Hong Kong's interest rates closely tracking those of the US, except for the peaks experienced during the Asian crisis of 1997-1998. 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). The turbulent period lasted until end-1998, as illustrated by the development in the interest rates in Figure 3.

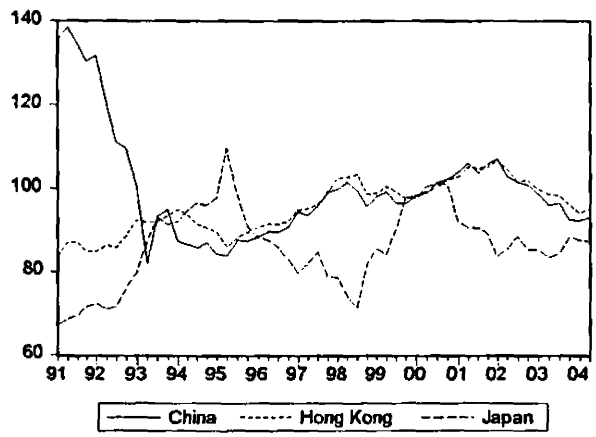


Figure 2. Nominal effective exchange rates (index, foreign currency price of home currency). China, Hong Kong and Japan, 1991-2004.

Chinese consumer prices started to fall in 1998 as the economy slowed down with the emergence of the Asian financial crisis. Deflation lasted until 2000 and reappeared again in late 2001. Finally, positive annual inflation rates prevailed in early 2003. A rather limited number of studies have focused on the recent inflation dynamics in this economy. Ha *et al.* (2003) found that the value of the renminbi, world prices and the level of productivity were important determinants of long-term price movements in China. During the era of low inflation and deflation, higher productivity and an appreciation of the effective exchange rate were found to be important explanatory variables for prices. 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, where a long appreciation has given way to the recent depreciation. The latter has served to intensify complaints about an undervalued renminbi, especially from the US. The IMF (2003) argued that transitory (lower commodity prices and WTO-related tariff cuts) and long-term (productivity gains, new technology and increased competition) supply shocks were behind the recent deflation in China. In similar vein, Cargill and Parker (2004) argued that Chinese deflation was supply-led. Gradual deregulation is taking place in the conduct of monetary policy, as Chinese banks are able to depart by increasing ranges from the reference rates on credits and deposits. However, as pointed out by Korhonen

(2004), Chinese interest rates have been nominally rather stable and high given the deflationary climate of the recent years. The stable behaviour is evident in the behaviour of the average repurchase rate for China since mid-1999, shown in Figure 3 (again, the series only begins in 1996 due to reasons of data availability).

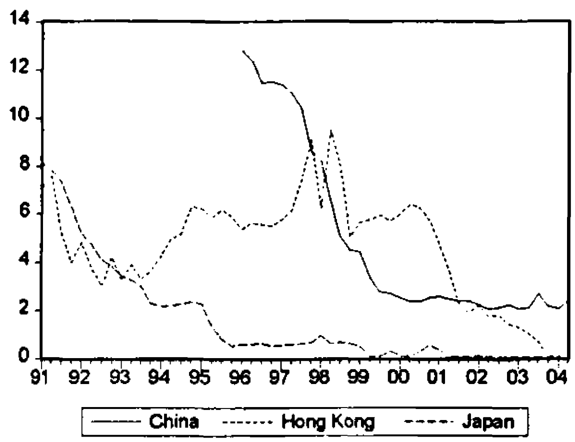


Figure 3. Short-term interest rates (3-month rate for Japan and Hong Kong, average repurchase rate for China), 1991-2004.

4.3 Theoretical Considerations

In this section, we discuss some theoretical considerations relevant to our research question. The experience of the Japanese deflation and the prolonged economic stagnation in that economy have given rise to a number of studies suggesting monetary policy strategies to avoid the zero interest rate floor or ways to escape from it once "trapped". While the latter ones are possibly of more relevance to our work, 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.

Perhaps the most prominent suggestion of escape from the liquidity trap has been the 'foolproof way' proposed by Svensson (2001, 2003). This proposal consisted of three elements, as follows (here we follow the presentation by Svensson, 2001). First, the central bank should announce an upward-sloping target path $\left\{\hat{p}_t\right\}_{t=t_0}^{\infty}$ for the domestic price level, the latter defined as the (log) prices of domestic goods p_t:

$$\hat{p}_t = \hat{p}_{t_0} + \hat{\pi}(t - t_0), \quad t \geq t_0 \tag{4.1}$$

Note that hats denote target levels, stars signify foreign variables and π_t is (log) domestic inflation in period t defined as $\pi_t \equiv p_t - p_{t-1}$. Additionally, we define the (log) exchange rate s_t in units of domestic currency per unit of foreign currency, with the (log) real exchange rate amounting to $q_t \equiv s_t + p_t^* - p_t$.[3] Here the price-level target for the current period t_0 exceeds the current price level:

$$\hat{p}_{t_0} \succ p_{t_0} \tag{4.2}$$

A small and positive inflation target is additionally set, $\hat{\pi} \succ 0$.

In the second step, an announcement of a currency devaluation is made, together with a currency peg to a crawling exchange rate target,

$$s_t = \bar{s}_t, \quad t \geq t_0 \tag{4.3}$$

with the exchange rate target $\bar{s}_t$ given by

$$\bar{s}_t = \bar{s}_{t_0} + (\hat{\pi} - \pi^*)(t - t_0), t \geq t_0 \tag{4.4}$$

The central bank therefore commits to buying and selling unlimited amounts of foreign exchange at the exchange rate $\bar{s}_t$. The initial exchange rate target after the devaluation is chosen in order to ensure a real depreciation of the domestic currency relative to the steady state:

$$q_{t_0} \equiv p_{t_0}^* + \bar{s}_{t_0} - p_{t_0} \succ q \tag{4.5}$$

Because we assume the domestic price level to be sticky, the real exchange rate moves one-to-one with the nominal exchange rate in the short run. The exchange rate target corresponds to a nominal depreciation of the domestic currency at the rate of difference between the domestic inflation target and foreign inflation, $\hat{\pi} - \pi^*$. Nominal appreciation takes place in the case where $\hat{\pi} \prec \pi^*$; the peg is fixed instead of crawling if $\hat{\pi} = \pi^*$.

[3] This is consistent with the original presentation of the model. In contrast, we define the (log) index of the nominal effective exchange rate in the model estimations in units of foreign currency per unit of home currency.

Thirdly, an exit strategy is declared. When the price level target path has been reached, the currency peg will be abandoned in favour of flexible price level targeting, with the former target path or flexible inflation targeting with the previous 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 higher price level expectations in the private sector, having an impact on the real *ex ante* interest rate, and therefore stimulating the economy. This solves the credibility problem emphasized by Krugman (1998); it is difficult to make the private sector believe in higher inflation, especially if the central bank's reputation for achieving its objectives is not impeccable. Moreover, the impacts on credibility and price expectations are advantageous. This is particularly the case given that the theoretical literature of new open economy macroeconomics often predicts the exchange rate pass-through to be slow and rather small in magnitude. This outcome obtains for example under a pricing-to-market assumption, whereby import price-setters simply set prices equal to those of domestic goods. Exchange rates may then be 'disconnected' from other macro variables.

Indeed, some empirical evidence supports the view of a slow and partial exchange rate pass-through, especially to final goods prices. Kara and Nelson (2002) found changes in the exchange rate and CPI inflation in the UK to be weakly connected, even if there was a strong relationship between import prices and the exchange rate. Similarly, based on cross-country evidence, Engel (2002) argued that there was a stronger (even if not a full) pass-through of exchange rates to imported goods prices than to final goods prices. This was also suggested by Burstein, Eichenbaum and Rebelo (2002) in their study of the behaviour of inflation after nine large post-1990 devaluations. The inflation rates were found to be low relative to the rate of devaluation. The authors argued that distribution costs, and substitution from imports to local goods, could account for the post-devaluation behaviour of prices.

McCallum (2000, 2003) suggested a policy rule for a central bank that has hit the zero bound, subordinating the nominal exchange rate rather than the interest rate to macroeconomic conditions:

$$\Delta s_t = \mu_0 + \mu_1(2 - \Delta p_t) + \mu_2(\bar{y}_t - y_t), \qquad \mu_1, \mu_2 \succ 0. \tag{4.6}$$

In the above, s_t represents the log of the home-country price of foreign exchange. According to the rule, the rate of depreciation Δs_t has to increase once inflation Δp_t and/or output y_t are below the target values (2 and $\bar{y}_t$, respectively, in the equation above), similarly to a conventional interest rate rule. Observing the relevant asset price almost continuously, open-market purchases would be conducted in order to attain the desired level of the exchange rate. Instead of relying on such a rule-based approach, Orphanides and Wieland (2000) focused on the effects of monetary base on aggregate demand and inflation even at the zero bound. Their calibrated open-economy model included both the interest and exchange rate channels of monetary transmission. The impact emphasized by the authors was derived from the portfolio balance effect, where the exchange rate would react to changes in the relative domestic and foreign money supplies.

Coenen and Wieland (2003) considered the effectiveness of the three aforementioned proposals by Orphanides and Wieland (2000), McCallum (2000, 2003) and Svensson (2001, 2003). This was done in the context of a macroeconomic model that featured rational expectations and nominal rigidities, and was estimated for the United States, the euro area and Japan. The authors found the beggar-thy-neighbour effects of the proposals to be non-negligible and that the zero lower bound could induce significant losses in terms of output and inflation stabilization in Japan.

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 possess favourable insulation properties with respect to foreign inflationary pressures, allowing for independent monetary policy (for an early contribution, see Friedman, 1953). However, the empirical evidence in favour of the insulation properties of flexible exchange rates is somewhat ambiguous. Ghosh *et al.* (1997) surprisingly found both the level and variability of inflation to be substantially lower for fixed exchange rate regimes. In contrast, Quirk (1994) concluded that there was no significant relationship between the exchange rate regime and inflation behaviour. Of course, for an economy with flexible exchange rates that has hit the zero interest bound, conventional interest rate policy is constrained and external deflationary pressures are difficult to counteract. This may increase the attractiveness of interventions in foreign exchange markets.[4]

Media discussion in the context of Chinese deflation has also centred on the possible transmission of falling prices in this economy to other economies

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

in the region, as well as to the United States. The IMF (2003) reported evidence concerning the impact of price fluctuations in China on other economies in the context of a vector autoregression analysis. Regarding the countries in our study, the IMF found there was a "discernible but small" impact on Japanese prices, but a stronger impact on Hong Kong SAR, owing to its strong links with the Mainland. 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 in favour of this possibility, however, as the importance of exchange rate shocks for movements in Chinese consumer prices is found to be relatively 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.[5] Benhabib, Schmitt-Grohé and Uribe (2002) showed how Taylor rules could drive an economy to a liquidity trap by self-fulfilling decelerating inflation paths and aggregate fluctuations. In the resulting state, the central bank is unable to pursue conventional interest rate policy to reverse a downward slide of prices. The authors demonstrated how the destabilizing properties of Taylor-type rules emerge by considering the zero bound on nominal interest rates and global equilibrium dynamics. Similarly, Kuttner and Posen (2004) discussed the difficulties that underlie policy assessments with Taylor-type rules for the Japanese economy. These stem primarily from two sources: the difficulty of ascertaining the level of potential output, and the role of expectations in the central bank policy (the latter being crucial when the zero bound is hit).

Measures have been considered to make the interest rate channel potent even at the zero bound. One approach would render the zero bound irrelevant altogether by making negative interest rates feasible. This could be achieved by means of a currency tax, as recently emphasised by Buiter and Panigirt-

[5]In the discussion above, a depreciation of the home currency is predominantly seen as a desired outcome for the deflationary economy. Posen (1998) took a different stance, arguing that Japan should avoid intentional yen depreciation due to the increased uncertainty and wealth erosion caused by such a policy. The author also mentioned that currency depreciation may easily become uncontrollable, and that the benefits would mainly accrue to Japan's exporting sector. Interestingly, we found that a yen depreciation would increase real GDP in the Japanese economy, whereas the outcome was the opposite in China; a depreciation in the renminbi nominal effective exchange rate was found to lead to a *fall* in output.

zoglou (2003) and Goodfriend (2000). Imposing a carrying cost on money, the central bank would be able to achieve any negative interest rate. Of course, the administrative costs of such a policy may be non-negligible. These arise from the fact that coin and currency are bearer bonds: as the holders are anonymous and ownership is sufficient for the bonds to be payable, it is difficult to make the bondholders actually pay the coupon for the issuer. Taking a different approach, Eggertsson and Woodford (2003) and Jung, Teranishi and Watanabe (2001) showed how optimal policy in the liquidity trap involves creating private sector expectations of higher future inflation, reducing real interest rates and stimulating aggregate demand. Not only short-term real interest rates were shown to have an impact on aggregate demand; rather, the entire path of expected short-term real rates (or long-term rates) matter. Thus, a policy defined in terms of the nominal interest rates can be stimulative for aggregate spending even at the lower bound, provided that the public expects the central bank to keep nominal interest rates low for a considerable period into the future and not to offset future inflation.

4.4 Methodology

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

4.4.1 About the Data

For all three economies, we used a four-variable system comprised of indicators for real output, consumer prices, interest rates, and the nominal effective exchange rate. Frequency of the variables varied depending on data availability. For Japan, we used quarterly data for the period 1991Q1-2004Q2, while the estimation sample for Hong Kong is longer by just one observation, reaching until 2004Q3. In contrast, monthly data was used for the Chinese economy for the period 1996:1-2004M8.[6] This choice of estimation sample allows for the inclusion of the disinflationary and deflationary periods for

[6] It could have been possible to use monthly data also in the case of Japan, since data on output is available for this economy in the form of industrial production. However, misspecification tests produced unsatisfactory results for long lag lengths; the latter

both Japan and Hong Kong, while the financial bubble in Japan during the late 1980s is excluded from the analysis. Therefore, we are able to determine the role of the interest and exchange rates in the slide of the two economies towards deflation, while the "inflationary" period of the 1980s is left unaccounted for. As the Chinese interest rate series only commences in 1996, it is preferable (or rather, imperative) to employ monthly data for this economy in order to have an adequate number of observations in the analysis. The data sources are the OECD Main Economic Indicators, the International Financial Statistics and Thomson Datastream databases. All output and price series in levels are seasonally adjusted. A logarithmic transformation was applied to all other series except interest rates, and the Chinese year-on-year inflation rate. The data sources and transformations are mentioned in greater detail in Appendix A. Estimations were conducted using the software JMulTi (2004), version 3.11. All the series used in the estimations are also depicted in Appendix A.

We consider here each of the system variables in turn. For real output, we used real GDP for Japan and Hong Kong. For China, the choice of the output variable was particularly problematic. Quarterly real GDP series is only available from the year 2000 onwards. Industrial production (year-on-year growth rates) series are available from 1994 onwards at the IFS database, but this series finishes in October 2002. We therefore decided to use data on quarterly nominal GDP, available for the entire data sample, and linearly interpolated the series into monthly observations. Finally, we constructed a consumer price index series (only annual inflation rates were available) in order to deflate the GDP series into a real output variable.[7] For prices, the consumer price index was used for Japan and Hong Kong. Annual CPI inflation rates were used in the case of China, as no official consumer price

are arguably necessary to adequately capture the dynamics between the variables when monthly data is used. Furthermore, the use of monthly exchange rate data in macroeconomic analysis may be regarded as containing too much noise; see the discussion by Ito (2005).

[7]Data on *cumulative* year-on-year GDP growth rates would also have been available for the whole data sample for China. Such an approach was used in an earlier version of the paper, but it made comparisons of the results between the three economies rather difficult. 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 to be employed (such as industrial production) to yield information about monthly movements in output for our interpolation. Moreover, the estimation sample would have been too short for only lower frequency data for China to be employed in the system. Finally, it is not clear that other methods of output data construction would have produced preferable GDP series, given arguments concerning the quality problems of official Chinese GDP data, see e.g. Rawski (2001).

index was available. The use of this variable is also econometrically preferable 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. We can then interpret the NEER as capturing the effects of external conditions on the deflation episodes in these economies. More generally, the variable can be considered an indicator of a country's international competitiveness in terms of its foreign exchange rate, calculated by using the value of a country's trade with the trading partner countries and areas as its weights (for detailed formulae, see e.g. Bank of Japan, 2003). 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 the 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. 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, with the emergence of the zero interest rate policy in Japan 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 was still fluctuating, albeit at very low levels, until early 2001. In the case of China, the one year lending rate that serves as an important reference rate is administratively set. Consequently, only eight changes in its level are displayed by the data during our estimation period, 1996:01-2004:08.[8] Therefore, we have decided to use the average repurchase rate that is found to rather closely track the administratively set one year lending rate. Figure 4 displays the reference lending rate and the repurchase rate used in our study. The coefficient for contemporaneous correlation between the series is found to be high, amounting to 0.98.

[8] Despite the low frequency of changes in the nominal interest rate, the PBoC did react to the slowdown of 1996-1998, attempting to boost investment by lowering administrated nominal interest rates five times over this period (Roberts and Tyers, 2003).

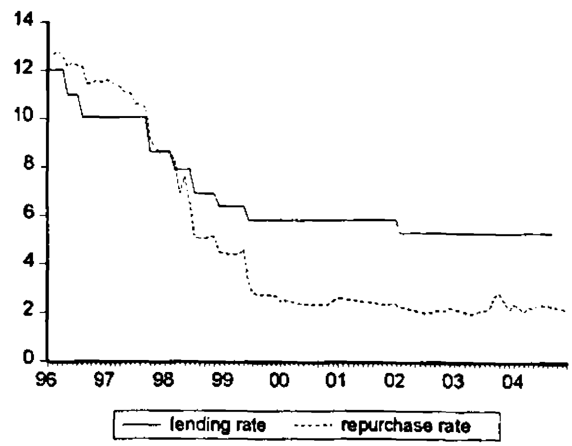


Figure 4. One-year reference rate for lending and average repurchase rate. China, 1996-2004.

In order to determine the order of integration of the underlying series, Augmented Dickey-Fuller (ADF) tests were performed for all the series. A Pantula principle was utilized in the testing procedure (see Pantula, 1989), whereby the series were initially differenced a sufficient number of times in order to obtain stationarity. Thus, the tests were initially conducted for all the series in second differences. Complete results from the unit root tests are listed in Appendix B. For Japan, Hong Kong and China, the nominal interest rate and the nominal effective exchange rate were found to be integrated of order one, I(1), so that first differencing would be necessary in order to render the series stationary. The only exception was the interest rate series for Japan where the ADF test provided evidence that the series in levels were stationary; however, an alternative test for unit roots, the KPSS test proposed by Kwiatkowski *et al.* (1992) suggested 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 were used in the test. Similarly, for all economies, the series for real GDP was found to be integrated of order one at a 5% level. Testing on the price series suggested that the price level for Japan and the annual inflation rate for China were integrated of order one.[9] In the case of Hong Kong, a unit root test with a structural break gave evidence that the price series in first differences was stationary. This result is attained by including a shift dummy in the system, obtaining a value of zero before

[9]The Japanese price level was 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.

1998Q3 and one thereafter. This date corresponds to the Asian crisis and the rapid drop in the rate of inflation at that time. We continued with the assumption that all the series in levels, together with the annual inflation rate for China, were 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.

The impact of exchange or interest rates on prices could be analyzed in a single equation framework, for example by estimating a simple AS-curve (such as a New Keynesian Phillips curve where the nominal effective exchange rate would enter the real marginal cost variable). However, a vector autoregressive (VAR) approach would allow for the simultaneous examination of interest rate and exchange rate shocks. 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, where 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 the exogenous shocks.

As the underlying series were found to be integrated of order one, it is possible that the series share common stochastic trends, i.e. cointegrating relationships. Testing for such relations in our very short sample for Hong Kong and Japan is not likely to yield reliable estimates, however.[10] Moreover, if the system was 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. 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 where there are limited signs of the economies experiencing cycles of any kind - the slowdown in the Japanese economy for a major part of the estimation sample and the enduringly high growth rates for the Chinese economy serve as prime examples.

[10] If the maximum lag length in the Saikkonen-Lütkepohl test (see Saikkonen and Lütkepohl, 2000) is set to 4, thus taking into account the extremely short sample size, a cointegration rank of zero could not 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 would suggest the existence of a cointegration rank of one for this economy.

Our choice, then, was to estimate the model in levels for Japan and Hong Kong, while in the case of China we opted for estimating a model in first differences, as such a procedure yielded more satisfactory results in terms of the misspecification tests. Admittedly, the differences between the model specifications make comparisons of the results problematic. In our view, analysing cointegration (long-run) relationships in an economy that is undergoing deep structural change is rather unappealing. Moreover, when the Saikkonen-Lütkepohl cointegration test (see Saikkonen and Lütkepohl, 2000) was used in the testing for common stochastic trends for China, the null hypothesis of no cointegration could not be rejected at conventional significance levels for any lag length suggested by the various information criteria, including a constant and trend as the deterministic terms. The results were consistent when applying the tests on trivariate subsystems, as shown in Appendix C.

Taking all the previous considerations into account, we opted for the estimation of a structural VAR model with contemporaneous restrictions, ignoring the possible cointegration relations. Such an approach is relatively common in the literature. Cushman and Zha (1997) used a similar methodology for Canada, a small open economy under flexible exchange rates, whereas Kim (1999) examined the G-7 group of countries. In a later study, Kim and Roubini (2000) focused on the G-7 excluding the US. In contrast, Jang and Ogaki (2003) used long-run restrictions in a structural vector error correction (SVEC) model to study the impacts of Japanese monetary policy shocks on exchange rates. However, the responses of the system variables to exchange rate shocks were not reported in the paper, and many of the illustrated impulse responses to contractionary monetary policy shocks showed signs of a price puzzle. Moreover, none of the studies mentioned above examined 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, that of Jang and Ogaki (2003) in 1993, while Kim (1999) extends the analysis through 1996. Finally, to our knowledge, a structural VAR model with contemporaneous restrictions has not yet been used for Hong Kong or China in order to examine the effects of monetary policy during the deflationary period.

4.4.2 The SVAR Approach

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

$$x_t = A_1 x_{t-1} + ... + A_p x_{t-p} + CD_t + u_t \qquad (4.7)$$

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 (4.7) can be expressed as:

$$\mathrm{A}x_t = A_1^* x_{t-1} + ... + A_p^* x_{t-p} + \mathrm{B}\varepsilon_t \tag{4.8}$$

where $\varepsilon_t \sim (0, I_K)$. Here the matrix A allows for the modeling of the instantaneous relations. Again, 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 that here the deterministic terms are omitted since they are not affected by the impulses hitting the system, nor do they affect such impulses themselves. The structural shocks are not predictable with respect to the past of the process and are the input for generating the time-series vector x_t. Moreover, they are assumed to be mutually uncorrelated and therefore orthogonal. As it is not possible to directly observe these structural shocks, certain assumptions are necessary for identification. The connection between the reduced and structural forms is obtained simply by multiplying (4.8) 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 can then be expressed as in Breitung *et al.* (2004):

$$u_t = \mathrm{A}^{-1}\mathrm{B}\varepsilon_t \tag{4.9}$$

We estimate the so-called AB-model of Amisano and Giannini (1997). In this case, the model for the innovations can be written as $\mathrm{A}u_t = \mathrm{B}\varepsilon_t$. Linear restrictions on A are written in explicit form as $\mathrm{vec}(\mathrm{A}) = R_\mathrm{A}\gamma_\mathrm{A} + r_\mathrm{A}$ 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 $\mathrm{vec}(\mathrm{B}) = R_\mathrm{B}\gamma_\mathrm{B} + r_\mathrm{B}$. Together, these two sets of restrictions are used to identify the system, i.e. the matrices A and B. The number of nonredundant elements of Σ_u, $K(K+1)/2$, is the maximum number of identifiable parameters in the matrices A and B. The overall number of elements in the structural form matrices A and B is $2K^2$. We therefore need to impose $2K^2 - \frac{K(K+1)}{2}$ further restrictions to identify the full model. These are discussed in the following.

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 then output, CPI, monetary policy (interest rate), and exchange rate shocks, respectively. The AB-model in the form $Au_t = B\varepsilon_t$ is written in the case of Japan as

$$\begin{bmatrix} 1 & 0 & 0 & 0 \\ a_{21} & 1 & 0 & 0 \\ 0 & a_{32} & 1 & 0 \\ a_{41} & a_{42} & a_{43} & 1 \end{bmatrix} \begin{bmatrix} u_t^y \\ u_t^p \\ u_t^i \\ u_t^{neer} \end{bmatrix} = \begin{bmatrix} b_{11} & 0 & 0 & 0 \\ 0 & b_{22} & 0 & 0 \\ 0 & 0 & b_{33} & 0 \\ 0 & 0 & 0 & b_{44} \end{bmatrix} \begin{bmatrix} \varepsilon_t^y \\ \varepsilon_t^p \\ \varepsilon_t^i \\ \varepsilon_t^{neer} \end{bmatrix} \quad (4.10)$$

As $2K^2 - \frac{K(K+1)}{2}$ restrictions are needed for exact identification (22 in our case of four endogenous variables), our model is over-identified. 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. Moreover, the non-contemporaneous impact of the exchange rate on prices is 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, where the interest rate is set on the basis of the price level only, not contemporaneously reacting to the exchange rate. As argued by Smets (1997), such behaviour is characteristic of an interest rate-targeting regime that largely neglects movements in the exchange rate in its policy. Moreover, it could well characterize the behaviour of the monetary authority in a relatively closed economy with flexible exchange rates such as Japan.[12] Finally, all variables are allowed to have a contemporaneous impact on the exchange rate. The exchange and interest rates, together with the price level, are all expressed in terms of quarterly averages as in the data source - this also ensures the appropriateness of our identification scheme.

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

[12] The small trade-to-GDP ratio of the Japanese economy is not consistent with the large interventions carried out by the monetary authorities in 2003-2004, as pointed out by Ito (2005). Possible exchange rate stabilization by the BOJ is further discussed in the following paragraphs.

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. Firstly, one could argue against the use of the price level in favour 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. One could also claim that once the inflation rate has reached negative territory, the central bank would be more interested in the effects of policy on changes in the price level than in the inflation rate, with the aim of escaping from the deflationary trap.[13] Regarding the timing assumption, Kim and Roubini (2000) and Cushman and Zha (1997) assumed that the central bank does not have 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; when using quarterly data as in our paper, 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 the 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, in order 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 outright 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 also that due to the contemporaneous nature of the identification scheme, no restrictions are imposed on the lags of the effects.

Despite the relatively closed nature of the Japanese economy and the flexible exchange rate arrangement, an argument could be made of the BOJ following a type of semi-exchange rate peg, where the exchange rate would significantly enter the objective function of the monetary authority. In particular, the unprecedented large (7% of GDP) foreign exchange interventions between January 2003 and March 2004 point to this possibility.[14] Of course,

[13] Some other studies of Japanese monetary policy are in effect using the price level instead of the inflation rate (see e.g. Morsink and Bayoumi, 2000; Shioji, 2000)

[14] Strictly speaking, the Ministry of Finance has the authority to intervene in Japan, with the Bank of Japan carrying out the interventions. Ito (2005) mentioned that the Bank of Japan had failed to affect the exchange rate in the desired direction between

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 has aimed at stabilizing the exchange rate. The authors found the estimated coefficient on the inflation rate in the BOJ reaction function to be less than one, casting doubt on the role of the rate of inflation as the target variable. However, the impact of a real exchange rate shock on the interest rate was not significant at a 5% level in the impulse response analysis. Contrasting evidence about the importance of the exchange rate in the monetary policy reaction function was also provided by Jinushi *et al.* (2000), who estimated such an equation for the pre-bubble period 1975-1985; in their specification, the exchange rate mostly obtained an insignificant coefficient. 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 to be an unrestricted element in the matrix, thereby not imposing it as zero. The resulting system is just-identified and can be considered a robustness test for the results obtained by estimating the benchmark model for Japan.

It is important to note that the identification scheme and the corresponding restrictions are based on the exchange rate arrangements and other considerations related to the monetary regimes of the three economies, not directly on the various theories regarding the use of the exchange rate channel to escape from a liquidity trap. Therefore, we are only able to comment on whether the results obtained provide support for 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 particularly the case due to the peg of the Hong Kong dollar and the Chinese renminbi to the US currency during our examination period. Of course, the currency peg may be loosened and a more active exchange rate management provided for. But if our results are used for the purpose of evaluating the feasibility of alternative monetary policy regimes, the Lucas critique may be of major importance - it is not clear that the behaviour of economic agents and thus our estimates, obtained in a backward looking scheme such as a vector autoregression, would remain robust to alternative policy arrangements.

The identification scheme for Hong Kong differs from the one of Japan 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

April 1991 and June 1995, whereas after that interventions became effective.

Hong Kong currency is set at 7.8 HKD per 1 USD. This type of arrangement of course qualifies as an extreme form of a currency peg. The monetary authority in a small open economy is likely to react quickly to the exchange rate shocks, reflected in strong interest rate movements.[15] We therefore leave the coefficient on the exchange rate a_{34} in the monetary policy reaction function unrestricted.[16] In contrast, 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, supporting the setting of their coefficients a_{31} and a_{32} in the interest rate equation to zero.[17] The model for Hong Kong is then written as

$$\begin{bmatrix} 1 & 0 & 0 & 0 \\ a_{21} & 1 & 0 & 0 \\ 0 & 0 & 1 & a_{34} \\ a_{41} & a_{42} & a_{43} & 1 \end{bmatrix} \begin{bmatrix} u_t^y \\ u_t^p \\ u_t^i \\ u_t^{neer} \end{bmatrix} = \begin{bmatrix} b_{11} & 0 & 0 & 0 \\ 0 & b_{22} & 0 & 0 \\ 0 & 0 & b_{33} & 0 \\ 0 & 0 & 0 & b_{44} \end{bmatrix} \begin{bmatrix} \varepsilon_t^y \\ \varepsilon_t^p \\ \varepsilon_t^i \\ \varepsilon_t^{neer} \end{bmatrix} \tag{4.11}$$

Finally, we discuss the identification scheme for China. Even if the currency peg to the US dollar would in ordinary circumstances lead to the characterization of the Chinese monetary authority as an exchange rate targeter, capital controls could in principle allow for the pursuit of an independent monetary policy.[18] Indeed, China's capital controls include restrictions on futures trading of the renminbi, foreign borrowing by Chinese enterprises,

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

[16]Admittedly, the nominal effective exchange rate used in our analysis is not pegged, reflecting the influence of external forces on Hong Kong's price level. However, the behaviour of the Hong Kong Monetary Authority certainly still qualifies as one of an exchange rate targeter, supporting our restriction.

[17]In contrast to the interest rate series for Japan, the Hong Kong and Chinese interest rates are defined as the rate prevailing in the end of the period, whereas the price level and exchange rate are period averages. While in the exchange rate equation it might have conceptually been preferable to use a period average for the interest rate also, 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.

[18]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 8.7 RMB = 1 USD, and the exchange rate was allowed to freely move within +- 0.25% of the previous day's reference rate (Huang and Wang, 2004). The exchange rate appreciated to 8.28 RMB = 1 USD by October 1997, where it remained until 2005. An intense public debate about the

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 within China. Korhonen (2004) noted that an increase in interest rates (a contractionary monetary policy shock) may have perverse effects in such an environment by encouraging saving and thus investment, further overheating the economy. Applying a closed economy assumption to China, similarly to Japan, we set the contemporaneous coefficient on the exchange rate in the monetary policy reaction function (a_{34}) to zero.[19] Note that while the closed economy assumption together with the regime of flexible exchange rates motivated the setting of the contemporaneous coefficient on exchange rates in the Japanese monetary policy reaction function to zero, an identical assumption is justified in the Chinese case by the existence of capital controls, despite the peg to US dollar. Furthermore, as we are using monthly data, availability constraints regarding output data (these are only available with a lag) justify the setting of the coefficient on output in the monetary policy reaction function (a_{31}) to zero, similarly to Japan. The over-identified model for China is then written identically to the one of Japan, system (4.10) above.

Despite the restrictions in the SVAR modelling 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 to identify 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 was 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 $u_t = B\varepsilon_t$.[20] When the B matrix is restricted to be lower triangular, the

possible undervaluation of the renminbi finally saw a revaluation of the Chinese currency by 2.1% against the US dollar in July 2005. At this time, the PBoC announced that the dollar peg was abandoned and the renminbi would be pegged to a basket of currencies.

[19]Examining the sources of real exchange rate fluctuations in China, Wang (2004) actually used an open-economy assumption, pointing to frequent movements in the nominal exchange rates that have reflected economic developments and the rates prevailing in the unofficial or swap markets. There is also some evidence that unsanctioned capital outflows have increased over time, as suggested by Yang and Tyers (2001).

[20]Naturally, a recursive identification scheme could also be expressed by imposing B to

first component of ε_t, ε_{1t}, can have an instantaneous impact in all four equations, whereas the second component ε_{2t} can have an effect on all 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 quite a close resemblance to the contemporaneous restrictions imposed in our model for other equations than the monetary policy reaction function. However, the recursive identification scheme has been defended in the context of reduced form VAR modelling 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 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). In contrast, the reduced form VAR can be estimated by simple OLS (Japan and Hong Kong) or feasible GLS (China, due to subset restrictions).

4.5 Estimation Results

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

The autoregressive order of the model was chosen on the basis of having a sufficient number of lags to satisfactorily capture the dynamic interaction between the variables, together with the results from the misspecification tests. An important additional consideration was the short sample size, possibly leading to a low estimation precision in the system including four endogenous variables. The models for Japan and Hong Kong were estimated using 4 lags, corresponding to one year of data, while 6 lags were included in the case of China where monthly data was used. Due to a high number of insignificant coefficients - commonplace in high-order VAR models - the model for China was reduced to a subset model by a procedure whereby the parameter with the lowest t-value in the model was checked and possibly eliminated from the system. Such a procedure is feasible in a system where all the variables are

be an identity matrix, with A lower triangular ($B = I_K$ and $Au_t = \varepsilon_t$).

stationary. Only coefficients with t-values above a threshold value of 1 were included in the final model for China. A constant and trend were included as deterministic terms for all the countries. The model for Japan included a shift dummy obtaining a value of one from 1997Q2 onwards and zero before. This dummy variable corresponds to a consumption tax hike in April 1997 that induced a shift in the consumer price index. We argue that it may be preferable to explicitly take this particular tax hike into account in our short sample where little movement in the price level is generally observable, making the shift in the series rather prominent. For Hong Kong, a shift dummy variable obtaining the value of 1 from 1998Q3 onwards was included. At this time, instability caused by the Asian crisis induced a rapid fall in the inflation rate.

Misspecification and stability tests for the reduced form subset VAR models are listed in Appendix D. We performed the Portmanteau and LM-tests for residual autocorrelation, the Jarque-Bera test for nonnormality and ARCH-LM tests to detect possible ARCH effects in model residuals. Neither of the two tests for autocorrelation, nor the one for ARCH effects suggest concern about model adequacy; only in the case of 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, the finding of nonnormality in the Jarque-Bera test may be of minor importance for our purposes. In the case of Japan, we additionally performed linearity tests as proposed by Teräsvirta (1994, 2004), where the linearity hypothesis was tested against a smooth transition regression model. These tests can be justified on the basis of the possible nonlinear behaviour of the interest rate near the zero floor, with the caveat that with integrated variables the validity of the tests is problematic. We conducted the tests on each of the four equations of the VAR model. In order to avoid singularity of the moment matrix, we set the deterministic terms to only appear linearly; similarly, whenever output, prices and the exchange rate were used as exogenous variables these were 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. Only for the price level equation did we find weak evidence of nonlinearity when the nominal interest rate at the fourth lag was used as the transition variable, but even then linearity could not be rejected at 1% level. In the case of the other three equations, we could not reject the null hypothesis of linearity for any of the four lags of our transition variable, justifying the use of our linear model.[21]

[21] These results are available from the author upon request.

No concern about the stability of our estimated model was revealed by the CUSUM test, based on the cumulative sum of model residuals, as 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. However, it may be prudent to take such evidence at a descriptive value only for Japan and Hong Kong; as our variables were found to be integrated of order one in unit root testing and appear in undifferenced form in the system for these two economies, the validity of the CUSUM-type tests is not certain (see Lütkepohl, 2004). Despite the short sample and a high lag order, the Chow forecast test can be used in order to evaluate the stability of the estimated systems, searching through the entire sample instead of focusing on some more arbitrary break dates that could be of interest for the analysis. Bootstrapped p-values are attractive in our case, since the approximate χ^2 and F-distributions of the Chow test statistics can be rather poor approximations and lead to very high rejection rates in small samples, as suggested by Candelon and Lütkepohl (2001). 1,000 bootstrapping replications were used. Stability of the estimated system was never rejected for Japan and China at a 5% significance level, while for Hong Kong smaller p-values were detected for 3 various break dates out of the total number of 27 tested break dates. As the majority of our evidence points to stable models, we proceeded with the estimation of the structural systems regarding the models as reasonably adequate for our analysis.

4.5.1 Japan

For Japan, the structural parameter estimates of the Λ matrix are written as follows. Note that we present the negation of the Λ matrix (the -Λ matrix), as interpretation of the contemporaneous coefficients is easier this way; a positive coefficient on the price level in the monetary policy reaction function then effectively corresponds to an increase in interest rates in response to a price hike, for example.

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

The asymptotic standard errors are reported in parentheses, 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. Whereas the point estimates of the contemporaneous impacts may not be of considerable interest, their signs could nevertheless be worthy of 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 ($-a_{42}$). A contractionary monetary policy shock similarly leads to a yen appreciation ($-a_{43}$), as expected. 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 of exchange rate depreciation in response to an output shock is less surprising though, if some of the recent fluctuations in the yen nominal exchange rate are considered. Even with the slowdown in the Japanese economy and the onset of deflation, the yen appreciated notably in 1995; similarly, a rapid acceleration in economic growth in 1996 was accompanied by exchange rate depreciation. Overall, we observe that the responses mostly obtain the expected signs but with quite limited statistical significance. More information could then be derived from a structural impulse response analysis. Taking advantage of the better small sample properties of bootstrap confidence intervals in comparison with other asymptotic methodologies, Hall bootstrap percentile 95% confidence intervals were used to illustrate parameter uncertainty. This particular approach additionally benefits from a built-in bias adjustment (see Benkwitz *et al.*, 2001). Responses up to 20 quarters ahead were considered, using 5,000 bootstrapping replications.

We mainly discuss here the responses to exchange rate and monetary policy (interest rate) shocks, as they are of primary interest to our analysis. The impacts of these shocks are depicted in Figure 5 below, with the monetary policy and exchange rate shocks 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, however, 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 would lower the price level by only 0.02 percent. This maximum impact is obtained after 5 quarters have passed from the shock. Interestingly, this estimate for the exchange rate pass-through is identical to the one found by Faruqee (2004) for the euro area - an economy with a similarly low degree of openness as that of Japan. Furthermore, our result is identical to the one by Gagnon and Ihrig (2004) for Japan during 1981-2003, while Ito *et al.* (2005) suggested an even lower coefficient (0.01). However, neither of these two estimates for Japan was statistically significant.

The impact of an exchange rate shock on the interest rate is negative in the short-run, then turning positive, but the impact is never statistically significant. The negative impact witnessed in the short run is expected, as an exchange rate appreciation is contractionary from a domestic policy viewpoint and would likely be met by a lower interest rate by the monetary authorities. However, the fact that this does not happen in the long run during our estimation sample could result from the zero interest rate floor and the limited movements in the interest rates caused by the zero bound. Interestingly, the yen has appreciated even in the middle of the Japanese recession - Bernanke (2000) argued that such movements suggest 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) worked through both a higher volume of exports and import prices, and by increasing the inflation expectations of the private sector. Our results indicate that all these channels are operative in the Japanese economy, with the 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 an exchange rate appreciation.

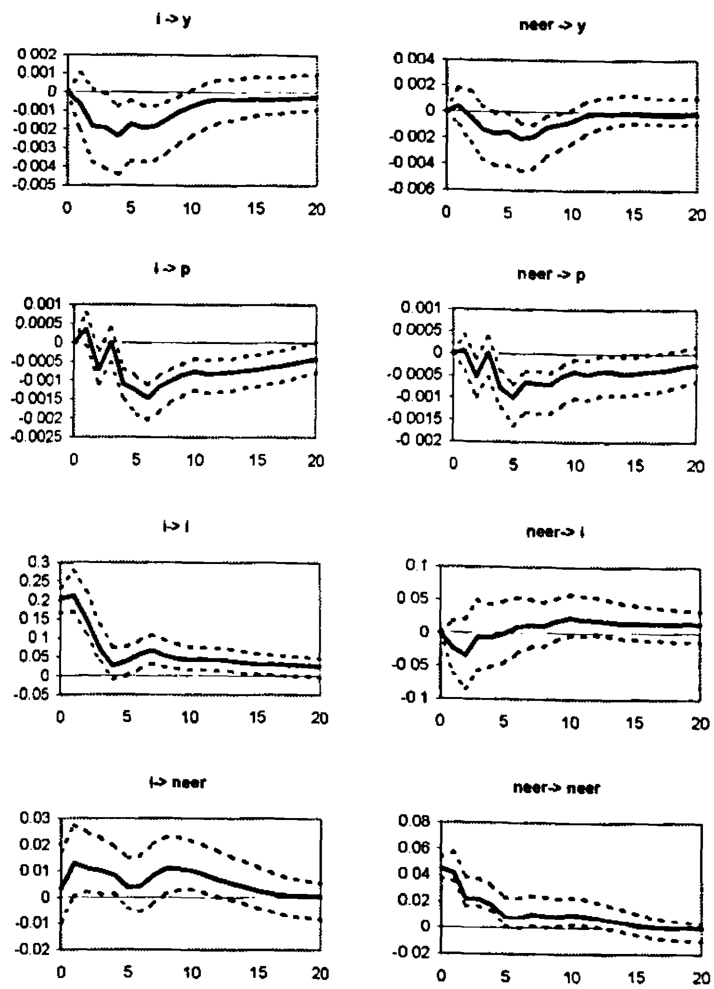


Figure 5. Impulse responses of output y, prices p, interest rate i and nominal effective exchange rate *neer* to a structural monetary policy shock (left column) and an exchange rate shock (right column), Japan.

As would be predicted by Keynesian models with price-wage inertia, but also 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 the 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.[22] 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. The real interest rate, which of course 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 leads to an exchange rate appreciation which is persistent and qualifies as delayed overshooting.

How would the results look if exchange rate stabilization was included as one of the objectives in the (contemporaneous) monetary policy reaction function for Japan, as described in the methodology section? Firstly, the signs on all the previous coefficient estimates in the structural parameter matrix A were found to remain robust. Further, the coefficient on the exchange rate ($-a_{34}$) in the interest rate equation obtained a negative (but statistically insignificant) sign; as expected, a contractionary movement in the exchange rate is met by a fall in the interest rate. The structural impulse responses to monetary policy and exchange rate shocks with 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. Using a 10% significance level instead, a statistically significant impact was obtained during 3 of the 20 quarters examined. Such a finding nevertheless suggests some caution when considering the feasibility of a strategy to use the exchange rate in order to escape from the liquidity trap.

In sum, both the interest rate and, to a less significant extent, the 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 took hold of the economy. The potency of the interest rate channel is interesting, as movements in the nominal interest rate have been rather limited and close to the zero lower bound. Our findings can be seen as

[22] Price puzzle generally refers to the response of the inflation rate rather than that of the price level; nevertheless, the response of prices in our system indicates that we have indeed identified a contractionary monetary policy shock.

supporting those of Ahearne *et al.* (2002) who examined the Japanese experience with deflation in order to draw implications for policy. The authors argued that there is little evidence of monetary policy having been ineffective in the early 1990s to ward off deflationary tendencies. Specifically, a more aggressive policy stance to fight deflation may have been beneficial still in 1993.[23] Indeed, in January 1993 - two years after the start of our estimation sample - the overnight uncollaterized call rate still stood at 3.9%, with annual inflation running at 1.2%. In similar vein, McCallum (2003) found that a comparison between the interest rates prescribed by a Taylor rule and the actual values for the overnight call rate make the BOJ's policy appear tight during 1993Q1-1998Q4. The finding of the potency of the interest rate channel would support a strategy to keep nominal interest rates at zero and committing to such strategy far into the future even when inflation eventually starts to increase, as suggested by Eggertsson and Woodford (2003) and Jung, Teranishi and Watanabe (2001). Moreover, additional stimulus could be implemented through negative interest rates by taxing currency, in line with arguments by Buiter and Panigirtzoglou (2003) and Goodfriend (2000). Such implications for the Japanese economy have been further discussed in a closed economy setup (without including an exchange rate channel) in our previous work (Mehrotra, 2005).[24]

Finally, it is important to note that the results can be seen as weakly supportive of a strategy of using the exchange rate channel in order to escape from the liquidity trap, in line with suggestions by Svensson (2001, 2003). This case, however, can only be made on the basis of the effects of the exchange rate shocks in our system, not on the basis of the structure of the model itself. In our system, exchange rate shocks could just arise in the foreign exchange market without being deliberately controlled by the central bank. It is also important to keep in mind that if the exchange rate was included in the (contemporaneous) central bank reaction function, the effect of the exchange rate on prices was only of borderline significance at a 5%

[23]The model simulations by Ahearne *et al.* (2002) suggested that if short-term rates had been lowered by another 200 basis points anytime between 1991 and early 1995, deflation could have been avoided.

[24]Monetary policy shocks are found to affect output and prices more quickly in our system including the exchange rate than in the closed-economy setup for Japan considered previously in Mehrotra (2005). This is likely to be the outcome of the different dynamics that the inclusion of a variable capturing (partly) external influences in the Japanese economy brings about. Indeed, we observe a similar outcome comparing the results of the two models (including and excluding the exchange rate) presented in Miyao (2000). The result could also be seen to suggest that the exchange rate channel is of importance for Japan and should be considered in the macroeconomic modelling of this economy.

level. But perhaps the biggest reason for caution is the low exchange rate pass-through to consumer prices suggested by our analysis. Moreover, the impact is spread through time. The strategy proposed by McCallum (2000, 2003) where 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 type of central bank reaction function. The nominal effective exchange rate in our model is an asset price that was allowed to react contemporaneously to new information in terms of output, prices and interest rates; to the extent that movements in the exchange rate reflected central bank interventions in the foreign exchange market, a McCallum-type rule could be implicitly supported by our results as well.

Our empirical results regarding the exchange rate shocks differ from those of Miyao (2000) who found that an appreciation shock (defined in terms of the nominal effective exchange rate - similarly to our model) led to an *increase* in Japanese output, while the system did not include a price variable of any kind. The study therefore did not provide evidence about the possibility of using the exchange rate channel to induce a positive change in the price level, even if the impact through the output channel did not seem to be operative in the direction suggested by the theoretical literature. Morsink and Bayoumi (2000) also reported an inclusion of the exchange rate in their VAR system to frequently yield perverse results, such as positive output response to an exchange rate appreciation. Kwon (1998) found a depreciation shock to the yen-dollar exchange rate to cause a fall in output, but 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), was found to lead to a fall in inflation - again contrasting with the feasibility to depreciate the yen in order to escape from the liquidity trap.[25]

4.5.2 Hong Kong

For Hong Kong, the structural parameter estimates for the A matrix are written as follows (consistently with the approach for Japan, we present the coefficients from the -A matrix to facilitate their interpretation).

[25] Of course, with limited controllability of the price level, the *real* exchange rate cannot be completely managed by the central bank. Thus, if the exchange rate is used as an instrument to fight deflation, the nominal exchange rate is the variable of interest.

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

The asymptotic standard errors are reported in parentheses. The likelihood ratio test does not reject the over-identifying 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 $-a_{21}$). 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 ($-a_{42}$). Therefore, most of the instantaneous impacts have the expected signs, even if one should note that independent monetary policy in this economy is restricted due to the currency board arrangement. Next, we proceed to examining the dynamics in the context of impulse response analysis. The effects in the system of exchange rate and monetary policy (interest rate) shocks are depicted below in Figure 6.

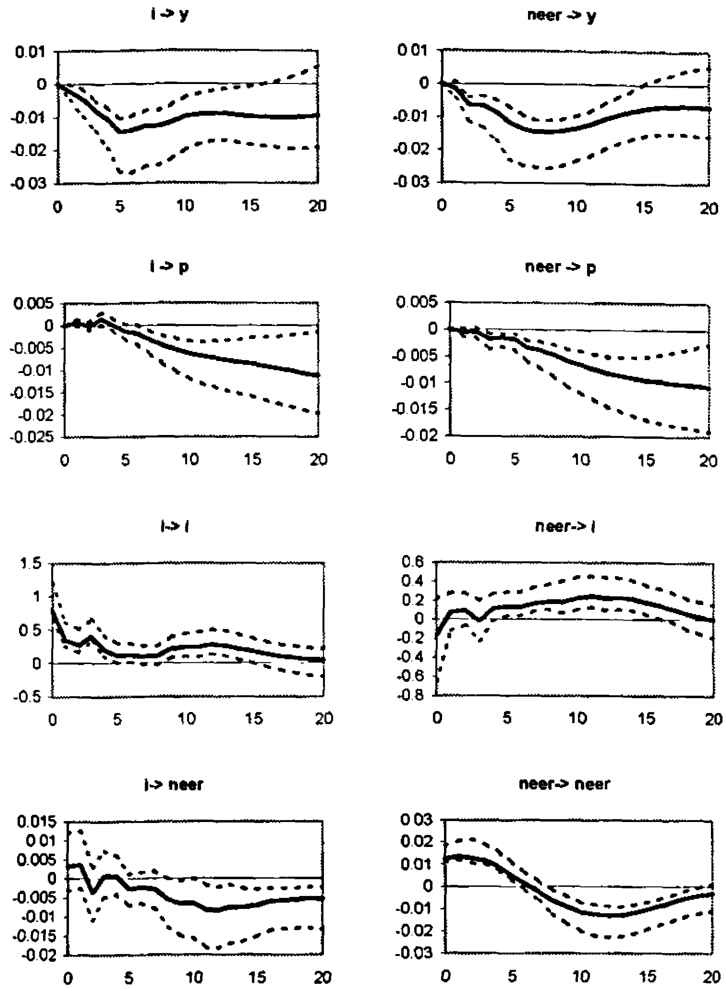


Figure 6. Impulse responses of output y, prices p, interest rate i and nominal effective exchange rate $neer$ to a structural monetary policy shock (left column) and an exchange rate shock (right column), Hong Kong.

An appreciation shock in the nominal effective exchange rate leads to a fall in real output, and similarly to the finding for Japan, the effect is significant. Importantly, an exchange rate shock also leads to a significant fall in the price level. It is of interest that the maximum point estimate of the pass-through to consumer prices during 20 quarters amounts to 0.77, significantly higher than in the case of Japan. The high pass-through is likely to result from the extremely open nature of the Hong Kong economy. Our finding is consistent with the result by Parsley (2003) who found a one-to-one pass-through from the nominal exchange rate to Hong Kong *import* prices. The effect of the exchange rate shock on the interest rate is negative in the short run but rapidly turns positive. Under a currency board, the interest rate adjusts to keep the exchange rate between the US and Hong Kong dollar constant and should not react to the effective exchange rate; it therefore appears difficult to explain why an effective exchange rate shock leads to a change in interest rates. However, as the Hong Kong interest rate follows the US rate, the fact that the nominal effective exchange rate of the Hong Kong dollar is entirely determined outside of Hong Kong may produce such a counterintuitive positive impact in 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, for we witness a significant decline in the price level in response to a monetary policy shock. This suggests that a monetary contraction also takes place in terms of the real interest rate. The increase in the return of domestic assets leads to an initial exchange rate appreciation but the exchange rate depreciates in the long run.

Here it should be emphasized that the exercise concerning the Hong Kong SAR is somewhat counterfactual, as the currency board arrangement does not make it possible for this economy to influence its nominal effective exchange rate in order to induce a positive change in the price level. In that sense, we are unable to comment directly on the desirability of a strategy to depreciate the Hong Kong dollar in order to counteract deflationary tendencies. However, the results should be seen in 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 usefulness of the currency board arrangement in an environment of falling prices.

In terms of the monetary policy shock, the fact that the responses of the price and output variables have the expected appearances does not mean, of course, that the policy implied by the US dollar peg would have been suitable for Hong Kong during the disinflationary and deflationary era. Rather, it is quite probable that an independent monetary policy would have responded more strongly to falling prices with the interest rate instrument. Evidence of this is provided in Table 1 below, in the form of annual inflation and the 3-month interest rates used in the study, for different points in time for both Japan and Hong Kong. Note that 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, with a rate of inflation of -5.29% one year later, the nominal interest rate still amounted to 5.94!

Time	Japan		Hong Kong	
	Inflation	3-month rate	Inflation	3-month rate
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 rates and 3-month rates of interest, Japan and Hong Kong

The contribution of the currency board to the deflationary episode has been somewhat downplayed by the Hong Kong Monetary Authority (2002). It was argued, however, that about half of the fall in the CPI can be attributed to a collapse of property prices that feed directly or indirectly to the CPI; it is also this part that is more likely to be influenced by domestic monetary conditions than cyclical factors, for example. Interestingly, Gerlach-Kristen (2004) used counterfactual simulations to study whether a Taylor-type interest rate rule or a reaction function in terms of the effective exchange rate would have produced different outcomes for the Hong Kong economy during 1990-2002. It was found that the inflation trajectories may have been different; deflation would have been less severe under a Taylor-rule, while economic output would have been hardly affected.[26]

4.5.3 China

For China, the structural parameter estimates for the A matrix are written as follows (in accordance with the approach for the two other economies, we present the coefficients from the -A matrix to facilitate their interpretation, with the asymptotic standard errors in parentheses):

[26] Lin and Lee (2002) examined the small open economies of Hong Kong and Taiwan (the latter with an exchange rate system of a managed float), and argued that macroeconomic performance was not systematically related to exchange rate regimes.

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

The likelihood ratio test for over-identifying restrictions does not suggest a rejection of our model, as 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 our 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 similarly to 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 very limited, such an outcome could well follow from the Chinese regime of capital controls, making the (purely external) exchange rate movements disconnected from the domestic economy.

More evidence about the dynamic interaction between the variables is obtained from impulse response analysis, conducted in an identical fashion to the two other economies. We depict the impulses to exchange rate and monetary policy (interest rate) shocks in Figure 7 below. As the underlying series in our estimated VAR system 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 and the small movements in the interest rates in China, especially considering the recent deflationary episode, it is quite remarkable that the responses to monetary policy shocks are consistent with conventional perceptions of such shocks - both output and the inflation rate fall. However, the impact on prices is only borderline significant at a 5% level. Similarly, the negative impact on output is no longer significant after 7 months have passed since the shock. Such a fast impact of a monetary policy shock on output in our system is somewhat worrying, even if the signs of the shocks provide evidence that we have indeed 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 enjoyed the status of prominent monetary policy instruments in the Chinese context. This arises even if the PBoC has been increasingly reliant on the adjustment of interest rates to achieve stabilization, as pointed out by Zhang and Wan (2002). The authors estimated Euler equations to examine the effects of Chinese monetary policy on the consumption of households, and nevertheless found inflation rates to be more relevant than nominal interest rates in determining consumption.

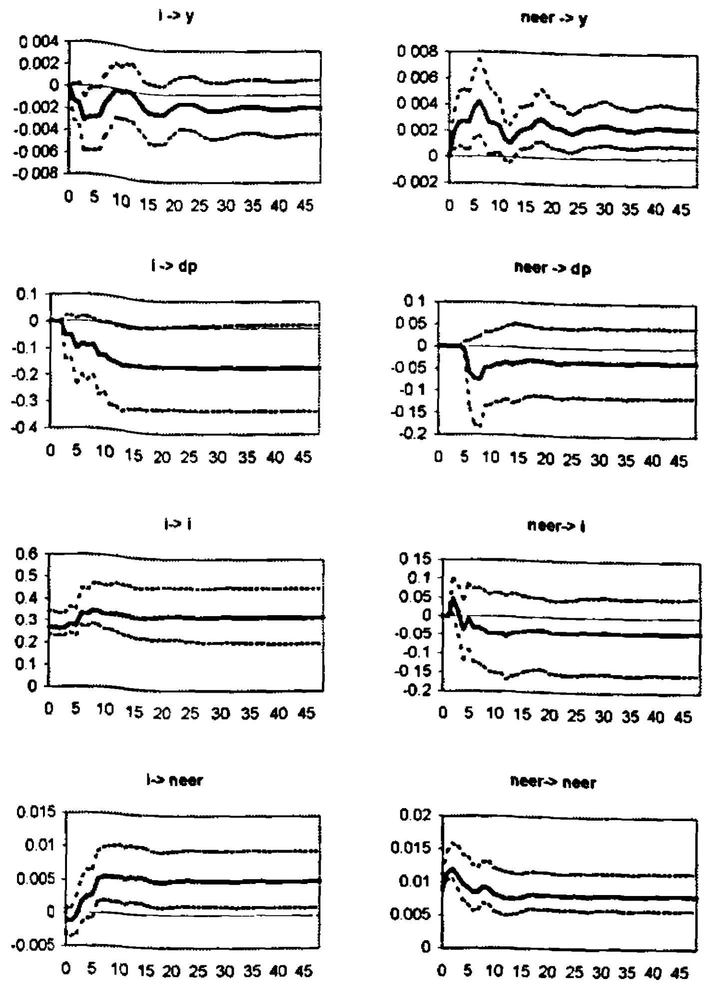


Figure 7. Impulse responses of output y, annual inflation rate Δp, interest rate i and nominal effective exchange rate $neer$ to a structural monetary policy shock (left column) and an exchange rate shock (right column), China.

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 when 8 months have passed from the shock). Moreover, the effect is not statistically significant. A counterintuitive finding is the effect of the exchange rate shock on output, as 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) could be offset by an accommodative monetary policy, for 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 - the annual inflation rate is indeed found to fall. However, we consider it 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 the vocal demands for Chinese renminbi appreciation most notably by the US - we find that such a measure may actually lead to an increase in the Chinese output, even if nothing is of course suggested about the impact on the Chinese-US trade balance.[27] Finally, the impact of a contractionary monetary policy shock on the nominal exchange rate is of interest. In the context of a theoretical model, Sun and Ma (2005) suggested 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 would be 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).

The degree of robustness for our results for Japan was already investigated by considering an alternative identification scheme where the monetary authority was assumed to contemporaneously react to the nominal effective exchange rate. For the two other economies, robustness tests were performed by considering a simple recursive identification scheme. Appendix F depicts the responses to monetary policy and exchange rate shocks, with the 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, as the nominal effective exchange rate is now found to depreciate following a positive shock to output. The coefficient on output in the monetary policy reaction function is found to be positive, while that on the price level is counterintuitively negative. This is likely to reflect the fact that the Hong Kong interest rates are determined endogenously without considering the local economic environment in terms of output or price developments. Of course, the Hong Kong business cycles may be correlated with those of the US - this would in effect explain the positive coefficient ob-

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

tained 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. This is not surprising, as the benchmark system for this economy is already recursive with one over-identifying restriction. 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.

4.5.4 Forecast Error Variance Decomposition for the Three Economies

In order to examine the importance of different shocks on the price level (inflation rate in the case of China), we used the forecast error variance decomposition (see Breitung *et al.*, 2004). The procedure calculates the contribution of one variable to the forecast error variance of another variable h periods ahead. As we are mainly 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.

Country	Horizon	shock y	shock p (Δp for China)	shock i	shock $necr$
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 in the annual inflation rate (China).

Regarding Japan, we can see that monetary policy shocks, measured by shocks to the interest rate, are dominating the shocks in the consumer price index in the long run. The proportion of inflationary shocks in the forecast error variance of consumer prices is very high in the short run, but greatly declines over time. In contrast, the importance of exchange rate shocks remains relatively constant, being at its highest after two years have passed since the shock. However, little evidence exists to support the argument that deflation was predominantly being imported to the Japanese economy. For Hong Kong, the results resemble those of Japan. Both monetary policy and inflationary shocks are found to be quite significant contributors to movements in consumer prices, and similarly to Japan, the importance of the latter decreases over time. But notably, the importance of exchange rate shocks is high in the long run, being twice as high in magnitude as that in Japan after 5 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 very little importance, with little variance through time. This is in stark contrast to their role detected in the other two economies; particularly monetary policy shocks in Japan and exchange rate shocks in the case of Hong Kong.

What could explain the relatively high importance of the exchange rate shocks in Japan in the short run (until 2 years have passed since the shock)? Campa and Goldberg (2002) found 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, which is in line with the limited insulation properties of fixed exchange rate regimes against shocks to foreign inflation. In addition, whereas 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 weak statistical significance when an alternative identification scheme was used. The use of the exchange rate channel in Japan should then be seen in light of increasing inflation expectations and lowering the real *ex ante* interest rate, as in the theoretical zero bound literature, rather than 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) was found to be significantly lower in Japan than in Hong Kong. Since our estimates of exchange rate pass-through for all three countries are high only in the case of Hong Kong, the results could be taken to suggest that exchange rate depreciation is most conducive to inflation in small open economies.

Our results for Hong Kong are also broadly in line with findings by Genberg (2003): we confirmed the relatively high importance of the nominal effective exchange rate (determined completely outside Hong Kong) for the price level in the long run, but also found the domestic monetary conditions to play a relevant role, as witnessed by the importance of monetary policy shocks for price movements during the disinflationary and deflationary era. 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 the *nominal* interest rates though) could be seen as contrary evidence to such a claim.[28] Of course, due to increased eco-

[28]Similarly, Yip and Wang (2002) suggested that the flexibility of the Hong Kong economy - argued by the proponents of the currency board to be high - may not be sufficient,

nomic integration, price convergence between Hong Kong and the Mainland could be of importance to 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. If these shocks were represented by the "own" inflationary shocks of our forecast error variance decomposition, then their share in our system is rather small in the long run, amounting to only 10% after five years.

For China, another economy with an exchange rate peg, our results suggest a very limited importance of exchange rate shocks in the determination of consumer prices. This finding arises consistently with the low share of fuel imports and the large share of manufactures in China's import composition,[29] price regulations and, probably more importantly, the existence of capital controls that could in principle allow for an independent monetary policy. Our results are somewhat in contrast with those of Ha *et al.* (2003) who found that the value of the renminbi and world prices were important determinants of long-term price movements in China, with the former explanatory variable occupying a prominent role during the era of low inflation and deflation. But even if the nominal effective exchange rate does not appear to be a significant factor in our analysis, world prices could still be of importance - these may actually be captured by the inflationary shocks themselves in our system. These shocks are of (perhaps surprisingly) prominent importance during our estimation sample, for their share always accounts for over 90% of the overall shocks. The limited use of a conventional interest rate 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 could be of major importance.

4.6 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

taking into account the high experienced volatility in GDP and export volume growth.

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

economies: Japan, Hong Kong and China. We estimated an open economy structural vector autoregressive (SVAR) model with contemporaneous restrictions for the three economies with different exchange rate regimes and monetary policy arrangements. In Japan and Hong Kong, an appreciation shock in the nominal effective exchange rate leads to a statistically significant fall in the price level; however, the impact is found to be considerably higher in the latter economy. Similarly, in both these two economies, the impact of the interest rate shock on prices is rather strong. In contrast, neither exchange nor interest rate shocks significantly influence price developments in China. The limited importance of the latter shocks is hardly surprising, as the Chinese monetary policy has predominantly operated through other (administrative) measures than market-determined interest rates. Similarly, capital controls in this economy may have largely insulated the economy from foreign shocks originating in the exchange rate.

Our results suggest that an appreciation of the nominal effective exchange rate could represent a powerful external deflationary factor especially in a small open economy; alternatively, a depreciation of the currency could provide a way to escape from the liquidity trap when the use of the conventional interest rate channel is limited. The latter argument is in line with suggestions in the theoretical literature, perhaps most prominently made 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. It is interesting that the importance of monetary policy shocks for consumer price movements in Hong Kong and Japan is relatively high during our estimation period when short-term interest rates have been close to or at the zero bound. In such an environment, if 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, or through the expectations channel whereby the central bank commits to keeping interest rates low for a considerable period of time into the future, as suggested by Eggertsson and Woodford (2003), and Jung, Teranishi and Watanabe (2001).

In a situation where many economies with flexible exchange rates or adjustable pegs are in a liquidity trap, they cannot simultaneously depreciate against one another. This was recognized in the "foolproof way" suggested by Svensson (2001, 2003). However, in our case with the (unadjusted) dollar pegs of Hong Kong and China, Japan could indeed depreciate its currency against the dollar. Admittedly, more active exchange rate management in

Japan in order to induce a positive change in the price level might create more accusations of 'beggar-thy-neighbour' policies, especially as the required changes in the exchange rates would need to be very large. Finally, our results do not yield information about the transmission of monetary shocks (both in terms of the interest and exchange rates) from one economy to another, as the nominal effective exchange rate takes into account many 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 research.

Bibliography

[1] Ahearne, A., Gagnon, J., Haltimaier, J., Kamin, S., 2002. Preventing Deflation: Lessons from Japan's Experience in the 1990s. Board of Governors of the Federal Reserve System, International Finance Discussion Paper no. 729.

[2] Amisano, G., Giannini, C., 1997. Topics in Structural VAR Econometrics. Springer-Verlag, Berlin.

[3] Andrade, J. P., Divino, J. A., 2005. Monetary Policy of the Bank of Japan - Inflation Target versus Exchange Rate Target. Japan and the World Economy 17, 189-208.

[4] Bank of Japan, 2003. Financial and Economic Statistics Monthly. December 2003, No. 57. Research and Statistics Department, Bank of Japan.

[5] Benhabib, J., Schmitt-Grohé, S., Uribe, M., 2002. Avoiding Liquidity Traps. Journal of Political Economy 110 (3), 535-563.

[6] Benkwitz, A., Lütkepohl, H., Wolters, J., 2001. Comparison of Bootstrap Confidence Intervals for Impulse Responses of German Monetary Systems. Macroeconomic Dynamics 5, 81-100.

[7] Bernanke, B. S., 2000. Japanese Monetary Policy: A Case of Self-Induced Paralysis? In: Mikitani, R., Posen, A. S. (Eds.), Japan's Financial Crisis and Its Parallels to US Experience. Institute for International Economics, Washington DC.

[8] Breitung, J., Brüggemann, R., Lütkepohl, H., 2004. Structural Vector Autoregressive Modeling and Impulse Responses. In: Lütkepohl, H., Krätzig, M. (Eds.), Applied Time Series Econometrics. Cambridge University Press, Cambridge.

[9] Buiter, W. H., Panigirtzoglou, N., 2003. Overcoming the Zero Bound on Nominal Interest Rates with Negative Interest on Currency: Gesell's Solution. The Economic Journal 113 (October), 723-746.

[10] Burstein, A., Eichenbaum, M., Rebelo, S., 2002. Why Are Rates of Inflation So Low After Large Devaluations? NBER Working Paper no. 8748.

[11] Campa, J. M., Goldberg, L. S., 2002. Exchange Rate Pass-Through into Import Prices: A Macro or Micro Phenomenon? NBER Working Paper no. 8934.

[12] Candelon, B., Lütkepohl, H., 2001. On the Reliability of Chow-Type Tests for Parameter Constancy in Multivariate Dynamic Models. Economics Letters 73, 155-160.

[13] Cargill, T. F., Parker, E., 2004. Price Deflation, Money Demand, and Monetary Policy Discontinuity: A Comparative View of Japan, China, and the United States. North American Journal of Economics and Finance 15, 125-147.

[14] Christiano, L. J., Eichenbaum, M., 1992. Liquidity Effects and the Monetary Transmission Mechanism. American Economic Review 82 (2), 346-353.

[15] Coenen, G., Wieland, V., 2003. The Zero-Interest-Rate Bound and the Role of the Exchange Rate for Monetary Policy in Japan. Journal of Monetary Economics 50 (5), 1071-1101.

[16] Cushman, D. O., Zha, T., 1997. Identifying Monetary Policy in a Small Open Economy under Flexible Exchange Rates. Journal of Monetary Economics 39, 433-448.

[17] Dolado, J. J., Lütkepohl, H., 1996. Making Wald Tests Work for Cointegrated VAR Systems. Econometric Reviews 15, 369-386.

[18] Eggertsson, G., Woodford, M., 2003. The Zero Bound on Interest Rates and Optimal Monetary Policy. Brookings Papers on Economic Activity 2003 (1), 139-211.

[19] Eichenbaum, M., Evans, C., 1995. Some Empirical Evidence on the Effects of Monetary Policy Shocks on Exchange Rates. Quarterly Journal of Economics 110, 975-1010.

[20] Engel, C., 2002. The Responsiveness of Consumer Prices to Exchange Rates and the Implications for Exchange-Rate Policy: A Survey of a Few Recent New Open-Economy Macro Models. NBER Working Paper no. 8725.

[21] Faruqee, H., 2004. Exchange Rate Pass-Through in the Euro Area: The Role of Asymmetric Pricing Behavior. IMF Working Paper no. 04/14.

[22] Friedman, M., 1953. The Case for Flexible Exchange Rates. In: Friedman, M. (Ed.), Essays in Positive Economics. University of Chicago Press, Chicago.

[23] Friedman, M., Schwartz, A. J., 1963. A Monetary History of the United States 1867-1960. Princeton University Press, Princeton.

[24] Fujii, E., 2004. Exchange Rate Pass-Through in the Deflationary Japan: How Effective is the Yen's Depreciation for Fighting Deflation? CESifo Working Paper no. 1134.

[25] Gagnon, J. E., Ihrig, J., 2004. Monetary Policy and Exchange Rate Pass-Through. Board of Governors of the Federal Reserve System, mimeo.

[26] Genberg, H., 2003. Foreign versus Domestic Factors as Sources of Macroeconomic Fluctuations in Hong Kong. Hong Kong Institute for Monetary Research Working Paper no. 17/2003.

[27] Genberg, H., Pauwels, L., 2003. Inflation in Hong Kong, SAR - in Search of a Transmission Mechanism. Hong Kong Institute of Monetary Research Working Paper no. 1/2003.

[28] Gerlach, S., 2005. Monetary Operations by Hong Kong's Currency Board. Journal of Asian Economics 15, 1119-1135.

[29] Gerlach-Kristen, P., 2004. Internal and External Shocks in Hong Kong: Empirical Evidence and Policy Options. University of Hong Kong, mimeo.

[30] Ghosh, A. R., Gulde, A., Ostry, J. D., Wolf, H. C., 1997. Does the Nominal Exchange Rate Regime Matter? NBER Working Paper no. 5874.

[31] Goodfriend, M., 2000. Overcoming the Zero Bound on Interest Rate Policy. Journal of Money, Credit and Banking 32 (4) Part 2, 1007-1035.

[32] Ha, J., Fan, K., 2004. One Country, Two Prices? A Study of the Deflationary Effect of Price Convergence on Hong Kong. Asian Economic Journal 18 (4), 405-421.

[33] Ha, J., Fan, K., Shu, C., 2003. The Causes of Inflation and Deflation in Mainland China. Hong Kong Monetary Authority Quarterly Bulletin, September 2003, 23-31.

[34] Hong Kong Monetary Authority, 2002. Hong Kong's Currency Board Today: The Unexpected Challenge of Deflation. Quarterly Bulletin 8/2002, 48-53.

[35] Huang, H., Wang, S., 2004. Exchange Rate Regimes: China's Experience and Choices. China Economic Review 15, 336-342.

[36] IMF, 2003. Deflation: Determinants, Risks and Policy Options - Findings of an Interdepartmental Task Force. International Monetary Fund Publications.

[37] Ito, T., 2005. The Exchange Rate in the Japanese Economy: The Past, Puzzles, and Prospects. The Japanese Economic Review 56 (1), 1-38.

[38] Ito, T., Sasaki, Y. N., Sato, K., 2005. Pass-Through of Exchange Rate Changes and Macroeconomic Shocks to Domestic Inflation in East Asian Countries. RIETI Discussion Paper no. 05-E-020.

[39] Jang, K., Ogaki, M., 2003. The Effects of Japanese Monetary Policy Shocks on Exchange Rates: A Structural Vector Error Correction Model Approach. Bank of Japan Monetary and Economic Studies, February 2003.

[40] Jinushi, T., Kuroki, Y., Miyao, R., 2000. Monetary Policy in Japan Since the Late 1980s: Delayed Policy Actions and Some Explanations. In: Posen, A. S., Mikitani, R. (Eds.), Japan's Financial Crisis and Its Parallels to US Experience. Institute for International Economics, Washington DC.

[41] JMulTi, 2004. Analysing Multiple Time Series with JAVA. Chair of Econometrics, Humboldt University Berlin. Available at www.jmulti.com.

[42] Jung, T., Teranishi, Y., Watanabe, T., 2001. Zero Bound on Nominal Interest Rates and Optimal Monetary Policy. Kyoto Institute of Economic Research Discussion Paper no. 525.

[43] Kara, A., Nelson, E., 2002. The Exchange Rate and Inflation in the UK. External MPC Unit Discussion Paper 11, Bank of England.

[44] Kim, S., 1999. Do Monetary Shocks Matter in the G-7 Countries? Using Common Identifying Assumptions about Monetary Policy across Countries. Journal of International Economics 48, 387-412.

[45] Kim, S., Roubini, N., 2000. Exchange Rate Anomalies in the Industrial Countries: A Solution with a Structural VAR Approach. Journal of Monetary Economics 45, 561-586.

[46] Korhonen, T., 2004. Monetary Regulation in a Changing China. In: BOFIT China Review 2004. Bank of Finland BOFIT - Institute for Economies in Transition Publications.

[47] Krugman, P., 1998. It's Baaack! Japan's Slump and the Return of the Liquidity Trap. Brookings Papers on Economic Activity 1998 (2), 137-187.

[48] Kuttner, K. K., Posen, A. S., 2004. The Difficulty of Discerning What's too Tight: Taylor Rules and Japanese Monetary Policy. The North American Journal of Economics and Finance 15 (1), 53-74.

[49] Kwiatkowski, D., Phillips, P. C. B., Schmidt, P., Shin, Y., 1992. Testing the Null of Stationarity Against the Alternative of a Unit Root: How Sure Are We That the Economic Time Series Have a Unit Root? Journal of Econometrics 54, 159-178.

[50] Kwon, E., 1998. Monetary Policy, Land Prices, and Collateral Effects on Economic Fluctuations: Evidence from Japan. Journal of the Japanese and International Economies 12, 175-203.

[51] Lanne, M., Lütkepohl, H., Saikkonen, P., 2002. Comparison of Unit Root Tests for Time Series with Level Shifts. Journal of Time Series Analysis 23 (6), 667-685.

[52] Lin, K. S., Lee, H-Y., 2002. The Political Economy of Exchange Rate Regimes: Evidence from Hong Kong and Taiwan. Asian Economic Journal 16 (1), 1-15.

[53] Lütkepohl, H., 1991. Introduction to Multiple Time Series Analysis. Springer-Verlag, Berlin.

[54] Lütkepohl, H., 2004. Vector Autoregressive and Vector Error Correction Models. In: Lütkepohl, H., Krätzig, M. (Eds.), Applied Time Series Econometrics. Cambridge University Press, Cambridge.

[55] McCallum, B. T., 2000. Theoretical Analysis Regarding a Zero Lower Bound on Nominal Interest Rates. Journal of Money, Credit, and Banking 32 (November, Part 2), 870-904.

[56] McCallum, B. T., 2003. Japanese Monetary Policy, 1991-2001. Federal Reserve Bank of Richmond Economic Quarterly 89 (1), 1-31.

[57] Mehrotra, A., 2005. The Case for Price Level or Inflation Targeting - What Happened to Monetary Policy Effectiveness during the Japanese Disinflation? European University Institute, mimeo.

[58] Meredith, G., 1998. Monetary Policy: Summary of Staff Views. In: Aghevli, B. B., Bayoumi, T., Meredith, G. (Eds.), Structural Change in Japan: Macroeconomic Impact and Policy Challenges. International Monetary Fund, Washington DC.

[59] Miyao, R., 2000. The Role of Monetary Policy in Japan: A Break in the 1990s? Journal of the Japanese and International Economies 14, 366-384.

[60] Morsink, J., Bayoumi, T., 2000. Monetary Policy Transmission in Japan. In: Bayoumi, T., Collyns, C. (Eds.), Post-Bubble Blues: How Japan Responded to Asset Price Collapse. International Monetary Fund, Washington DC.

[61] Orphanides, A., Wieland, V., 2000. Efficient Monetary Policy Design near Price Stability. Journal of the Japanese and International Economies 14, 327-365.

[62] Pantula, S. G., 1989. Testing for Unit Roots in Time Series Data. Econometric Theory 5, 256-271.

[63] Parsley, D. C., 2003. Exchange Rate Pass-Through in a Small Open Economy: Panel Evidence from Hong Kong. International Journal of Finance and Economics 8, 99-107.

[64] Posen, A. S., 1998. Restoring Japan's Economic Growth. Institute for International Economics, Washington DC.

[65] Quirk, P. J., 1994. Fixed or Floating Exchange Rate Regimes: Does it Matter for Inflation? IMF Working Paper no. 4134.

[66] Rawski, T. G., 2001. What is Happening to China's GDP Statistics? China Economic Review 12, 347-354.

[67] Roberts, I., Tyers, R., 2003. China's Exchange Rate Policy: The Case for Greater Flexibility. Asian Economic Journal 17 (2), 155-184.

[68] Saikkonen, P., Lütkepohl, H., 2000. Testing for the Cointegrating Rank of a VAR Process with an Intercept. Econometric Theory 16 (3), 373-406.

[69] Schellekens, P., 2003. Deflation in Hong Kong SAR. IMF Working Paper no. 03/77.

[70] Shioji, E., 2000. Identifying Monetary Policy Shocks in Japan. Journal of the Japanese and International Economies 14, 22-42.

[71] Smets, F., 1997. Measuring Monetary Policy Shocks in France, Germany and Italy: the Role of the Exchange Rate. Swiss Journal of Economics and Statistics 133, 597-616.

[72] Sun, H., Ma, Y., 2005. Policy Strategies to Deal with Revaluation Pressures on the Renminbi. China Economic Review 16, 103-117.

[73] Svensson, L. E. O., 1999. How Should Monetary Policy Be Conducted in an Era of Price Stability? In: New Challenges for Monetary Policy, a Symposium Sponsored by the Federal Reserve Bank of Kansas City, 26-28 August 1999, 195-259.

[74] Svensson, L. E. O., 2001. The Zero Bound in an Open Economy: A Foolproof Way of Escaping from a Liquidity Trap. Bank of Japan Monetary and Economic Studies 19, 277-312.

[75] Svensson, L. E. O., 2003. Escaping from a Liquidity Trap and Deflation: The Foolproof Way and Others. Journal of Economic Perspectives 17 (4), 145-166.

[76] Taylor, J. B., 1993. Discretion versus Rules in Practice. Carnegie-Rochester Conference Series on Public Policy 39, 195-214.

[77] Teräsvirta, T., 1994. Specification, Estimation, and Evaluation of Smooth Transition Autoregressive Models. Journal of the American Statistical Association 89, 208-218.

[78] Teräsvirta, T., 2004. Smooth Transition Regression Modeling. In: Lütkepohl, H., Krätzig, M. (Eds.), Applied Time Series Econometrics. Cambridge University Press, Cambridge.

[79] Wang, T., 2004. China: Sources of Real Exchange Rate Fluctuations. IMF Working Paper no. 04/18.

[80] Yang, Y., Tyers, R., 2001. The Asian Crisis and Economic Change in China. The Japanese Economic Review 52 (4), 491-510.

[81] Yip, P. S. L., Wang, R. F., 2002. Is Price in Hong Kong That Flexible? Evidence from the Export Sector. Asian Economic Journal 16 (2), 193-208.

[82] Zhang, Y., Wan, G. H., 2002. Household Consumption and Monetary Policy in China. China Economic Review 13, 27-52.

4.7 Appendices

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 Chinese nominal GDP series, the Hong Kong consumer price index and the real GDP for Hong Kong were seasonally adjusted by a Census X-11 procedure by the author, whereas the consumer price index and real GDP for Japan were seasonally adjusted at the data source. The Chinese nominal GDP was linearly interpolated to monthly observations and deflated by the consumer price index in order 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, setting a value of 100 for January 1993. Monthly observations were subsequently calculated by using the monthly year-on-year growth rate on consumer prices.

Figures of Series Used in Estimation

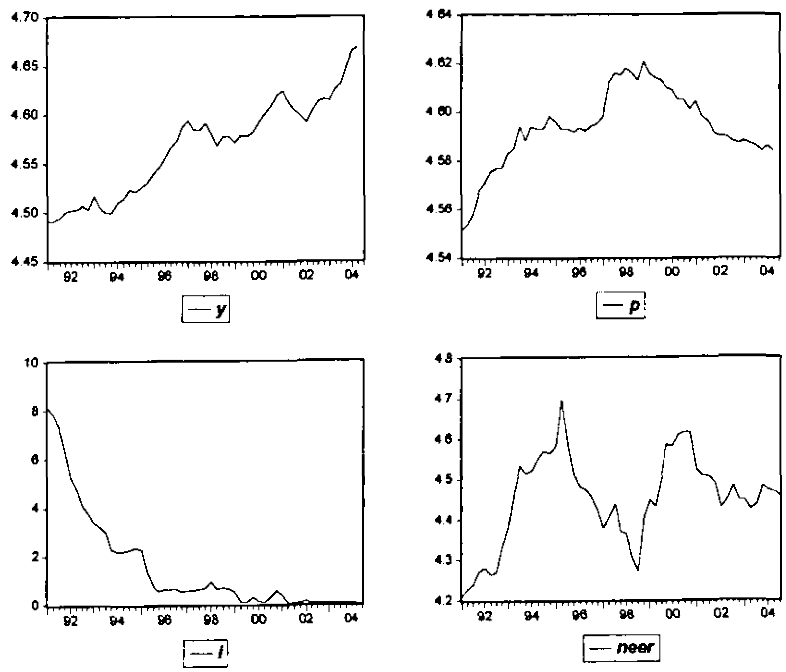


Figure: Series, Japan.

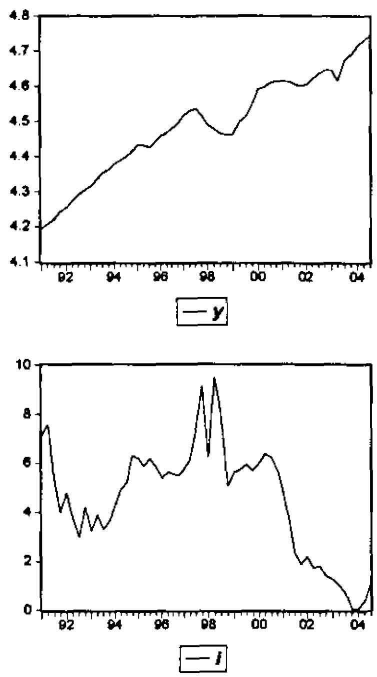


Figure: Series, Hong Kong.

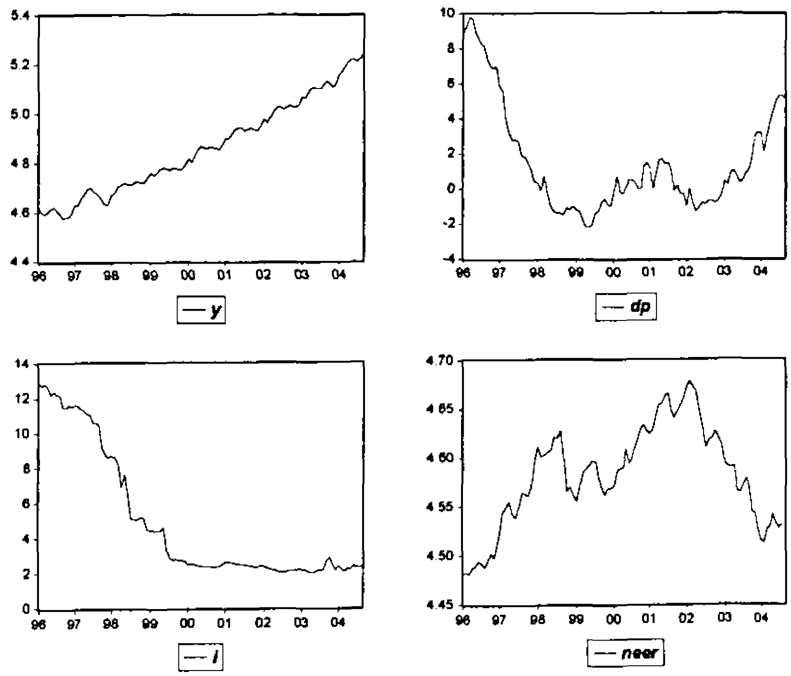


Figure: Series, China.

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

Note: Price series tested with unit root test with structural break (Lanne *et al.*, 2002)
Prefix S denotes date of shift dummy.
* indicates significance at 10% level, ** at 5% and *** at 1% level.
The order specification criteria in parentheses: AIC=Akaike, HQ=Hannan-Quinn, SC=Schwarz-criteria.
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)	-3.31*
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***
i	c, t	1 (AIC,HQ,SC)	-4.36***

Note: Price series additionally tested with unit root test with structural break (Lanne *et al.*, 2002).
Prefixes S and I denote the dates of the shift and impulse dummy variables, respectively.
* indicates significance at 10% level, ** at 5% and *** at 1% level.
The order specification criteria in parentheses: AIC=Akaike, HQ=Hannan-Quinn, SC=Schwarz-criteria.
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.
The order specification criteria in parentheses: 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	Det. term	No. of lags	Coint. rank	Test stat
$y, i, p, neer$	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.
The order specification criteria in parentheses (see Appendix B).

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

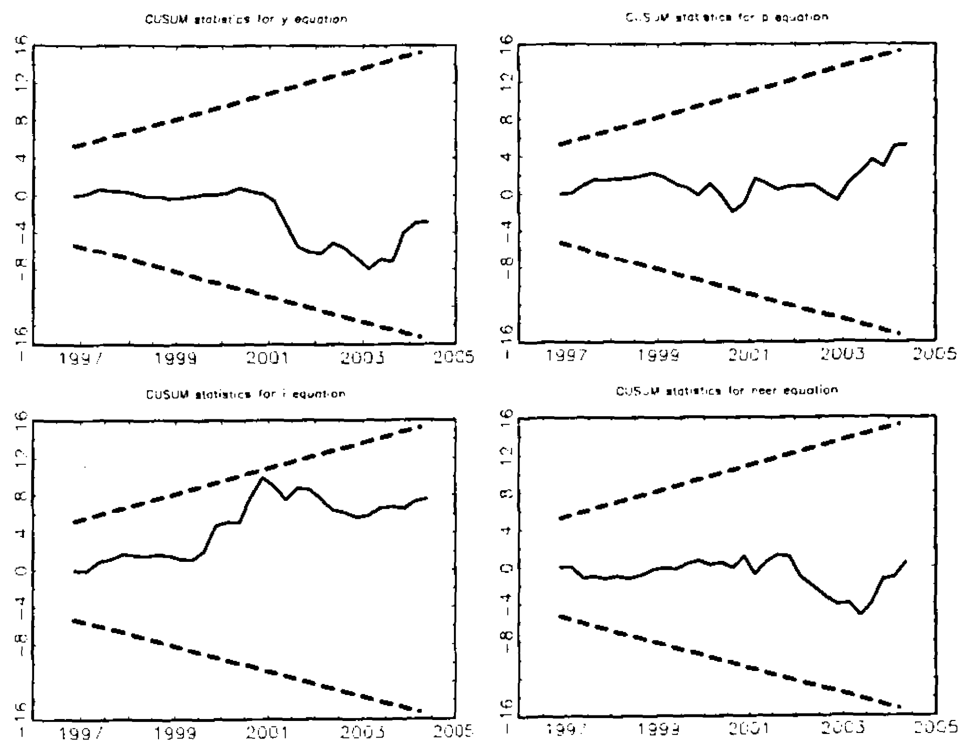
Q_{16}	179.97 [0.72]
FLM_5, FLM_4, FLM_1	1.42 [0.12], 1.23 [0.22], 1.54 [0.11]
$LJB(s_3^2), LJB(s_4^2)$	3.59 [0.46] 16.10 [0.00]
$ARCH_{LM}(16)$(eqs. 1, 2, 3, 4)	11.36 [0.79] 5.93 [0.99] 17.19 [0.37] 20.14 [0.21]

China

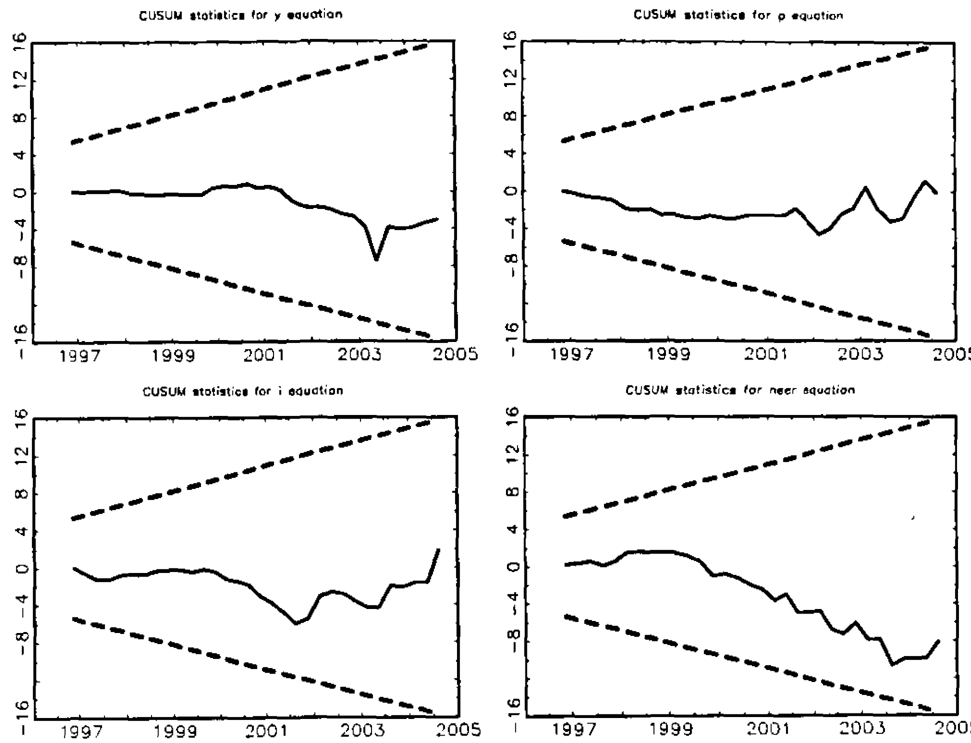
Q_{16}	203.11 [0.77]
LM_5, LM_4, LM_1	56.97 [0.98], 41.61 [0.99], 8.61 [0.93]
$LJB(s_3^2), LJB(s_4^2)$	21.94 [0.00] 17.66 [0.00]
$ARCH_{LM}(16)$(eqs. 1, 2, 3, 4)	18.11 [0.32] 7.33 [0.97] 42.74 [0.00] 11.24 [0.79]

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.

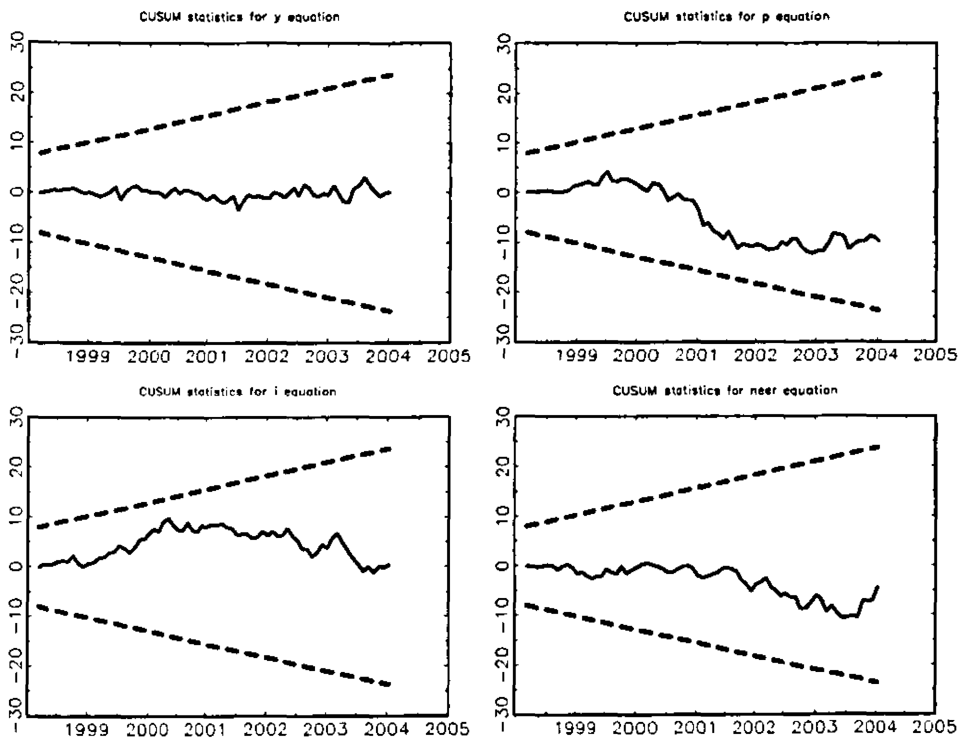
Stability Tests
CUSUM tests, 5% significance level



Japan


CUSUM statistics for y equation
CUSUM statistics for p equation
CUSUM statistics for i equation
CUSUM statistics for neer equation
16
12
8
4
0
-4
-8
-16
1997
1999
2001
2003
2005

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 parentheses)

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

Japan

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

Hong Kong

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

China

Appendix F

Structural Impulse Response Analysis, Alternative Identification Schemes

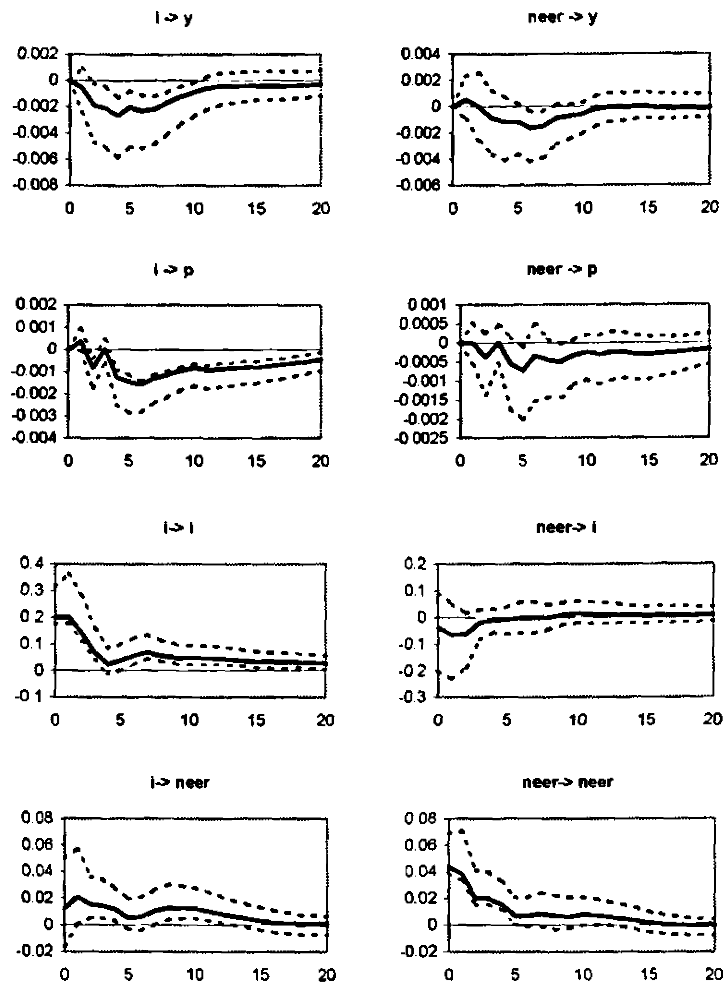


Figure: Impulse responses of output y, prices p, interest rate i and nominal effective exchange rate *neer* to a structural monetary policy shock (left column) and an exchange rate shock (right column); Japan, exchange rate targeting identification scheme.

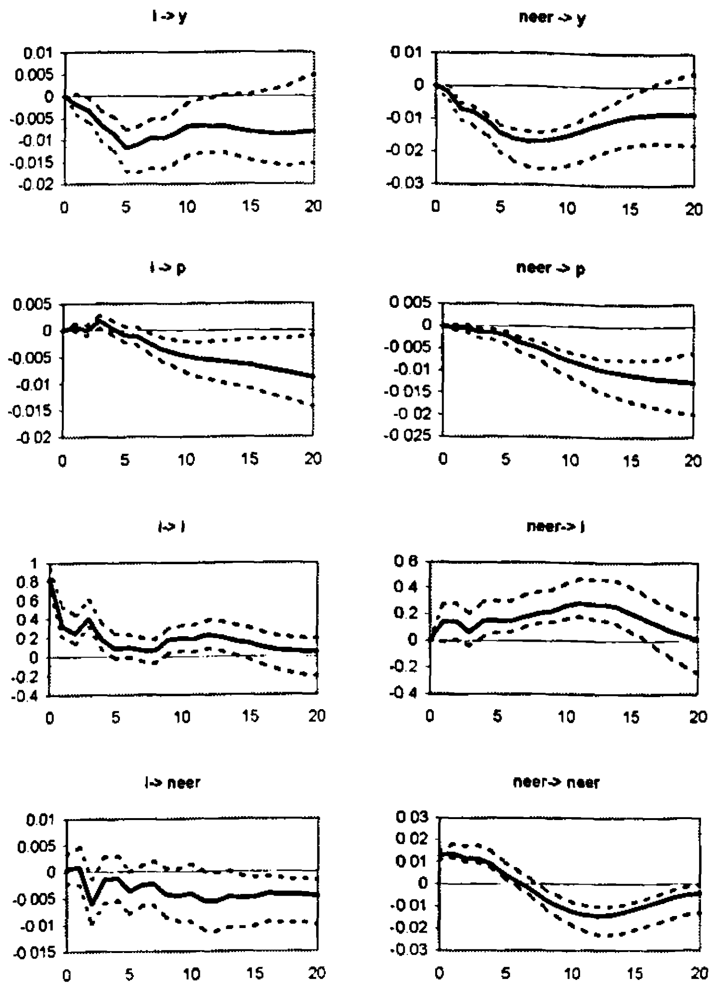


Figure: Impulse responses of output y, prices p, interest rate i and nominal effective exchange rate $neer$ to a structural monetary policy shock (left column) and an exchange rate shock (right column); Hong Kong, Cholesky decomposition.

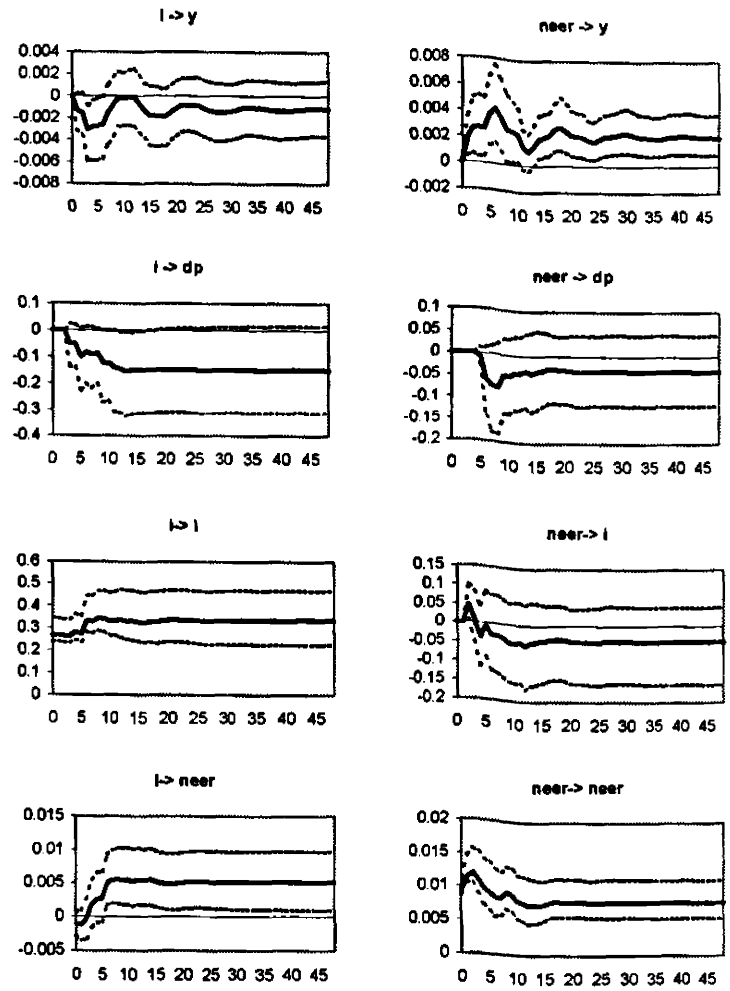


Figure: Impulse responses of output y, annual inflation rate Δp, interest rate i and nominal effective exchange rate *neer* to a structural monetary policy shock (left column) and an exchange rate shock (right column); China, Cholesky decomposition.

Chapter 5

Socio-Economic Development and Fiscal Policy - Lessons from the Cohesion Countries for the New Member States

Joint work with Tuomas Peltonen

5.1 Introduction

On 1 May 2004, the European Union (EU) expanded to a union of 25 Member States. An important feature of this enlargement is that many of the new Member States are still undergoing a transition process from command to market economies, and most of them fall quite far below the average EU income levels. Furthermore, the dispersion of income levels among these countries and their regions is striking (Vaitilingam, 2002). Despite the differences in their economic structures, the new Member States are expected to join the third stage of the European Monetary Union (EMU) and to follow considerable fiscal and monetary discipline prior to the adoption of the euro.

Fiscal policy will have an important role in the new Member States' economic policies during their process to adopt the euro. Upon EU accession, the new Member States are subject to the Treaty of Maastricht and the Stability and Growth Pact. According to a protocol to the Treaty, the general government deficit to GDP ratio should not exceed 3 percent and the public debt to GDP ratio should be lower than 60 percent. Additionally, the Stability and Growth Pact requires Member States to reach a budgetary position close to balance or in surplus over the medium term. However, the average fiscal deficit of the new Member States was 5.6 percent in 2003, with only the Baltic countries (Estonia, Lithuania and Latvia) and Slovenia fulfilling the excessive deficit threshold of 3 percent of GDP.[1]

Against this background, it is clear that if the new Member States aim at an early adoption of the euro, most of them need to consolidate their fiscal balances in order to meet the EU fiscal criteria. The implementation of fiscal consolidation might be a challenging task, as completion of the transition process together with the implementation of the *acquis communautaire* will increase the expenditure side of the government budgets, while pressures to introduce tax reforms could shrink the revenue side. Finally, the structural nature of the fiscal imbalances would certainly not ease this task.

The aim of our paper is to provide new insights into the convergence process of the new Member States[2] towards the common currency. More specifically, we want to examine the impact of different fiscal policy variables on socio-economic and structural development. We are especially interested in whether the fiscal austerity required by the Maastricht Treaty would restrict and be harmful for the socio-economic catching-up process of the new Member States. According to theory, if government spending and investment are efficient and beneficial for socio-economic development, fulfilling the fiscal criteria might be detrimental for the new Member States as many of them are currently running high government deficits and notable consolidation measures would be required. Alternatively, fiscal austerity could be beneficial for socio-economic development, for economic growth and stability and thus ultimately for welfare.

The socio-economic indicator that we calculate in our paper provides evidence that the Southern EU Member States, namely Portugal, Spain, and Greece, together with Ireland (cohesion countries henceforth) were at approximately the same level of socio-economic development in the 1980s, when they joined the European Union, as the new Member States were in 1999. Importantly, we also found that the levels of government debt and

[1] According to the Eurostat statistics. In contrast, the EU-15 fiscal deficit was 2.7 percent of GDP in 2003. However, the general government gross debt of the EU-15 amounted to 64.3 percent of GDP in 2003, while it was 42.1 percent in the new Member States.

[2] In this study the term new Member States is used to refer to the 10 new EU countries excluding Cyprus and Malta, but including the Accession Countries, Bulgaria and Romania. Note also that the term EU-15 is used for the Member States prior to the latest EU enlargement, excluding Luxembourg (for reasons of data availability). Furthermore, the term EU-11 is used to refer to the EU-15 less the cohesion countries.

net lending of the cohesion countries in the 1980s were highly similar to the respective variables in the new Member States in 1999.[3] This facilitates our analysis, as we can use data from these countries, where there is longer time series data available, to evaluate the relationship between socio-economic development and fiscal policy. Furthermore, the cohesion countries have also been subject to structural funds from the EU and capital inflows that are currently affecting the new Member States. Finally, the privatization of government enterprises started in the chosen time period in the cohesion countries, and this is also expected to happen in the new Member States.[4]

In the empirical part of the paper, we assess the extent to which the new Member States will have to adjust in order to achieve the socio-economic development levels of the former EU-15 Member States, and especially, what is the role of fiscal policy in order to achieve this aim. Our 'Socio-Economic Development Index' (SEDI) consists of various socio-economic indicators that are to a large extent affected by public policies. We measure the change in the development index during 1980-1999 in the cohesion countries and estimate the role of fiscal policy in the adjustment process. Using instrumental variable methods, we regress the SEDI on various fiscal variables, such as government primary surplus and public debt. Furthermore, we replicate the analysis for the other EU-11 Member States, in order to find whether fiscal consolidation would be equally relevant in promoting socio-economic development in the other EU countries. Finally, we evaluate the time it would take for the new Member States to achieve the EU benchmark levels in terms of the development indicator, assuming the average speed of development of the cohesion countries during 1980-1999. The times vary from 8.5 years (Slovenia) to around 24 years (Romania). However, it is important to note that the aforementioned analysis and its implications should not be confused with the convergence criteria that are a prerequisite for euro area entry.

Our results show that fiscal consolidation would be beneficial for socio-economic development in the medium term. In line with previous literature about the effects of fiscal consolidation on economic output, we find that fiscal retrenchment, including a lower level of public debt, would be advantageous to socio-economic development in the cohesion countries. The effects of fiscal consolidation are found to be more prominent in promoting socio-economic development in the cohesion countries than in the other EU-15

[3] Pelkmans *et al.* (2000) also suggested that this was the case for the level of GDP.

[4] There are also important differences. A significant one is that the new Member States are former command economies while the cohesion countries were market economies. This has an impact on the role and size of the public sector, and therefore on the level of socio-economic development.

Member States. Because the levels of socio-economic development, government debt and net lending of the new Member States in 1999 bear close resemblance to those of the cohesion countries in 1980s, they would seem to be the most relevant ones for our analysis. Findings from the transition literature suggest that those Central Eastern European countries that have adopted tighter fiscal policies in their transition process have been more successful with their stabilization policies and have experienced a faster recovery in output growth (see e.g. Budina and van Wijnbergen, 1997). These results emphasize the need for fiscal consolidation, in accordance with the Maastricht convergence criteria and additional recommendations from the EU Commission. They could also be seen as support for the Stability and Growth Pact or an equivalent intergovernmental device to curb public spending and debt. As a policy implication, new Member States wishing to increase their level of socio-economic development should pursue fiscal consolidation and pay attention to their government debt levels.

Next, we turn our focus to the relevant theoretical and empirical literature. As has been shown by e.g. Modigliani (1961), an increase in public debt decreases the capital stock of the economy (crowding out effect) and therefore lowers the growth rate of the economy. Furthermore, as shown by Diamond (1965) and later by Saint-Paul (1992), an increase in the level of public debt generally decreases the welfare of the economy. Moreover, a number of empirical studies have shown that fiscal retrenchment might have a favourable impact on economic activity in the medium term.[5] As explained by theoretical models, these non-Keynesian effects can occur through demand-side (effects on expectations, lowered risk premium, wealth effects) and supply-side channels (e.g. through increased competitiveness). Interestingly, according to the European Commission (2003), roughly half of the fiscal consolidation episodes undertaken in EU countries in the past three decades have been followed by an immediate acceleration in economic growth. In addition, the European Commission reports that fiscal consolidation has a positive impact on output in the medium term if it is conducted through expenditure retrenchment rather than through tax increases. Furthermore, Perotti (1999) found that fiscal consolidations are more likely to have non-Keynesian effects in countries with high debt levels.

In a conceptually similar study to ours, Afonso *et al.* (2003) examined public sector performance and efficiency in 23 OECD countries. The authors

[5]This line of research includes studies by Giavazzi and Pagano (1990, 1996), Alesina and Perotti (1995), Alesina and Ardagna (1998), Perotti (1999, 2002) and Giavazzi *et al.* (2000).

considered indicators for the 'opportunity-providing' activities of the government, such as education, health and infrastructure; and the 'Musgravian' tasks, such as allocation, distribution and stabilization. Whereas we consider a breakdown of fiscal balances, the study by Afonso *et al.* (2003) used total government spending in order to evaluate the level of public sector efficiency. Interestingly, the authors found that, when fiscal consolidation took place between 1990 and 2000, there was a considerable improvement in the public sector performance of the countries used in our study: Greece, Portugal, Spain and Ireland.

Concerning the link between structural reforms and fiscal policy, most of the transition literature sees the issue as a trade-off between structural reforms and fiscal balances, where rapid structural reforms may generate costs in the form of deteriorating fiscal balances (Pirttilä, 2001). This conclusion is drawn from the theoretical models such as Dewatripont and Roland (1992), Chadha and Coricelli (1997), and Coricelli (1998). If the transition process is seen as the release of factors of production from a declining state sector to an expanding private sector, as in Chadha and Coricelli (1997), then at least three factors contribute to the deterioration of the government budget balance. Firstly, the decline of the state sector decreases the established tax base. Secondly, the creation of a new and effective private sector tax system takes time. Thirdly, if there are frictions in the economy, the transition process is likely to result in higher unemployment, increasing the expenditures for unemployment benefits. However, the literature investigating the interaction of fiscal policy and socio-economic development is limited, and even more so, as far as the new Member States are concerned. Our paper tries to fill this gap, and provides policy recommendations for the new Member States in their convergence process.

The paper is structured as follows. The next section presents the empirical analysis, where we discuss the calculation of the Socio-Economic Development Index, our model specifications and the time-series properties of the data. This is followed by the estimation results, together with the possible implications for the new Member States. The final section concludes.

5.2 Empirical Analysis

Our main aim is to investigate the relationship between fiscal policy and socio-economic development in the four cohesion countries: Greece, Ireland, Spain and Portugal. In addition, we calculate the average speed of socio-economic development in these countries and then use this information to project the time required for the new Member States to attain similar levels of socio-economic development to the EU-15 Member States. Furthermore, we use the EU-11 Member States as a control group in the analysis about the relationship between socio-economic development and fiscal policy. The aim is to determine whether there are differences in the effects of fiscal consolidation in the promotion of socio-economic development between the two different groups of countries.

As mentioned before, the cohesion countries are relevant for our analysis, since their socio-economic development level and fiscal balances in the 1980s, at the time when they joined the EU, were highly similar to the ones of the new Member States in 1999. The estimation period, 1980-1999, captures the catching-up and the economic convergence period from the EU membership to the start of the third stage of the EMU.[6] Our estimation period was also characterized by capital flows from the EU to the respective economies, which were included in their government revenues, and are assumed to be approximately of the same magnitude as those of the new Member States.[7] In addition, the countries in the estimation sample experienced the privatization of government enterprises,[8] which is assumed to continue in the new Member States. Finally, our main assumption is that socio-economic development acts as an input and a catalyst for economic growth and convergence.

This section consists of four sub-sections. First, the Socio-Economic Development Index (SEDI) is derived. Second, the model specification issues are discussed. Third, the data sources and time series properties of the data are described. Finally, the evolution of the main variables is analyzed.

[6]Ireland became an EU member in 1973, Greece in 1981, and Spain and Portugal in 1986. As Ireland became an EU member in 1973, it may have been preferable to use an estimation period also covering the 1970s. However, data availability for the Socio-Economic Development Index for the period 1970-1979 is limited, and this period may be too early to capture the effects of EU accession in Portugal and Spain that only joined in 1986.

[7]Indeed, the EU budget foresees net flows to the new Member States to total 2.5%-4% of their gross national income during 2007-2013. This alleviation of the fiscal burden is comparable to that of the cohesion countries in the past, with the exception of Spain where the share was lower.

[8]The revenues from the sales of mobile phone licenses in the four countries in question are not relevant, because they are outside the sample period. The revenues were included in the government balances of 2000, 2001 or 2002, depending on the country in question.

5.2.1 Calculation of the SEDI

Assessment of the development level of the individual countries is based on the Socio-Economic Development Index (SEDI) that we derive in this section. The index consists of different indicators of health, infrastructure, environment and education. The SEDI is constructed to be as comprehensive an indicator of the level of socio-economic development as possible, taking into account the public/private sector nature of the variables and data limitations.

The variables that are included in the Socio-Economic Development Index are listed in Table 1 below. The data source for these variables is the World Bank World Development Indicators (WDI) 2003 Database.

Variables
Air passengers carried (per capita)
Railway passenger kilometers (1000km, per capita)
Telephone main lines in use (per 100 inhabitants)
GDP per unit of energy use (PPP USD per kg of oil equivalent)
Carbondioxide emissions (kg per 1995 USD GDP)
Primary school enrollment (% of gross population)
Tertiary school enrollment (% of gross population)
Infant mortality rate (per 1000 live births)
Immunisation DPT (% of children under 12 months)

Table 1. Variables in the Socio-Economic Development Index (SEDI).

In the SEDI, infrastructure is represented by the number of air passengers and railway passenger kilometers, as well as telephone main lines in use. The environmental variables used include carbondioxide emissions and the amount of GDP attained per unit of energy use. Both the primary and tertiary school enrollment are indicators of education and, finally, the level of public health is represented by the infant mortality rate and the rate of DPT immunisation. As argued by Afonso *et al.* (2003), these types of variables could be called 'opportunity' indicators, as a well-functioning health and education system provide many accessible opportunities for the population. As the authors claimed, the variables could also be seen as indicators of allocative efficiency.

The calculation of our index follows quite closely the one of the Human Development Index (HDI) of the United Nations (UN). There is, however, one major qualitative difference between the two indicators: unlike the UN index, our development indicator does not include the GDP level of the country in question.[9] One reason for this is that we regress the development index on a set of fiscal variables and the 'opportunity' indicators of our index are variables predominantly determined by government measures. Another reason is that GDP may not properly illustrate the welfare of the population. In the case of Ireland, for example, Laski and Römisch (2003) mentioned that there is a large difference between the GDP and GNP figures, and suggest that GNP may serve as a better measure of welfare. When net factor income from abroad is negative, as has increasingly been the case in Ireland in the 1990s, funds cannot be consumed nor saved in the country itself. Finally, we are controlling the GDP level on the right-hand side of the equation and having GDP on both sides of the equation might cause us some econometric problems.

The SEDI is calculated as follows.[10] First, we look for the smallest (min) and largest (max) absolute value for each variable j in the sample of 24 countries (EU-15 together with the new Member States excluding Luxembourg, Cyprus and Malta, but including Romania and Bulgaria) i for the period of 1980-1999.[11] In the case where a smaller value for a variable would correspond to a higher level of socio-economic and structural development, as is the case with the infant mortality rate and carbondioxide emissions, we use the inverse of the original values.[12] Then, the index number for any given observation (var in the formula below) for variable j for country i is yielded by:

$$index_{ij} = (var_{ij} - \min_j)/(\max_j - \min_j). \tag{5.1}$$

[9] The United Nations HDI measures a country's performance in terms of three different aspects of human development: longevity, knowledge and a decent standard of living. Longevity is measured by life expectancy at birth, knowledge by a combination of adult literacy rate and school enrollment at different levels. The standard of living is measured by GDP per capita. (United Nations, 2003)

[10] Another possibility would have been to proceed using the methodology of the UN in calculating the human development index, where 'goalposts' are selected, such as a maximum value of 85 and a minimum value of 25 for life expectancy. However, our methodology is very similar in that also the UN index has as its goalposts the feasible values at the extremes.

[11] However, in the case of the new Member States and Germany, we only used data from 1992 onwards in the construction of the SEDI. This was due to many missing variables for the new Member States before the start of the transition, and to the German unification that may have caused problems in the analysis.

[12] One could claim that a threshold level of emissions is necessary for a certain level of development, such as in the transition process from an agricultural to an industrial economy. However, industrialization had already taken place in the acceding countries, with heavy industries and excessive pollution being common phenomena.

From this construction it follows that all the values for $index_{ij}$ are between 0 and 1. One should note that as the values are obtained linearly, we implicitly assume that the fiscal measures would need to be as large as to get from value 0.1 to 0.2 as from 0.9 to 1.0. Therefore, we are assuming constant returns to scale, which is admittedly a constraining hypothesis.[13] The Socio-Economic Development Index for each country i is obtained by an arithmetic average of the $J = 9$ indices for country i.[14]

$$SEDI_i = \frac{\sum_{j=1}^{J} index_{ij}}{J}. \tag{5.2}$$

Next, we display a table with the ranks of the UN HDI for 2001 and our SEDI index for 1999, together with their values[15] (The UN index is predominantly based on data from 1999 and is thus comparable).

[13]Rzonca and Cizkowicz (2003) mentioned problems with using indices whose values are bounded at the extremes in econometric analysis. Our values for the countries under study, even if bounded by 0 and 1, fall in the middle of this range, with no visible slowdown in the growth rate of the index.

[14]The UN development index is also constructed by a simple average of the different 'dimension' indices: life expectancy, education and GDP. However, the weights within the dimensions vary. For example, in the education dimension, a 2/3 weight is given to adult literacy and a 1/3 weight to gross enrolment. (United Nations, 2003)

[15]In the UN HDI, we only list the countries included in our study.

SEDI	Value	HDI	Value
Denmark	0.766	Sweden	0.936
Sweden	0.748	Belgium	0.935
Ireland	0.733	Netherlands	0.931
Netherlands	0.715	Finland	0.925
France	0.706	France	0.924
Austria	0.699	United Kingdom	0.923
Finland	0.673	Denmark	0.921
Italy	0.667	Austria	0.921
United Kingdom	0.663	Germany	0.921
Germany	0.661	Ireland	0.916
Portugal	0.654	Italy	0.909
Belgium	0.645	Spain	0.908
Spain	0.640	Greece	0.881
Greece	0.592	Portugal	0.874
Slovenia	0.558	Slovenia	0.874
Hungary	0.549	Czech Republic	0.844
Czech Republic	0.494	Slovakia	0.831
Latvia	0.487	Hungary	0.829
Poland	0.476	Poland	0.828
Slovakia	0.450	Estonia	0.812
Lithuania	0.438	Lithuania	0.803
Estonia	0.425	Latvia	0.791
Bulgaria	0.387	Bulgaria	0.772
Romania	0.332	Romania	0.772
Average EU15	0.683		0.916
Average NMS+AC	0.460		0.816
(NMS+AC)/EU15	67.3 %		89.0 %

Table 2. Ranking and values of countries' SEDI and HDI. Sources: Authors' calculations and United Nations (2001).

As is clear from the previous table, the ranking in our Socio-Economic Development Index is strikingly close to the UN development index for the year 1999. The top performer in the UN index, Sweden, ranks second in ours. The two worst performers, Bulgaria and Romania, are the same in both indices. The biggest differences between the two indices are witnessed for Belgium, Denmark and Ireland.[16] The development levels of most of the new Member States are not far apart in our index. Slovenia at 0.558 and Hungary at 0.549 were in 1999 rather close to the level of Greece at 0.592. A country group consisting of the Czech Republic, Latvia and Poland were very close to one another. Slovakia, Lithuania and Estonia precede Bulgaria and Romania, which ranked lowest according to our index. As a comparison of the differences among the EU-15 Member States, Denmark's index was at 0.766, Germany at 0.661 and the lowest, as already mentioned, was Greece

[16]The SEDI and UN HDI indices differ in their composition, which explains the different ranking given by each index to the same country. For example, the very high scores in health and education explain the high position of Belgium in the UN HDI, whereas the country obtains somewhat lower values for the infrastructure and environmental variables that are included in the SEDI.

at 0.592. The EU-15 average in 1999 was 0.683. As a comparison the Socio-Economic Development Indicator is presented in Table 3 for the cohesion countries in 1980, 1999, and in the year the respective country joined the European Union, with the exception of Ireland that joined the EU in 1973. As can be seen from Table 3, at the time the cohesion countries joined the EU, they were at a comparable level of development to most of the new Member States in 1999.

SEDI	1980	1981	1986	1999
Greece	0.366	0.412		0.592
Ireland	0.308			0.733
Portugal	0.388		0.475	0.654
Spain	0.441		0.502	0.640

Table 3. Socio-Economic Development Index of the cohesion countries in 1980, 1999, and in the year the respective country joined the EU. Source: Authors' calculations.

Referring to the literature about the effects of fiscal consolidation on output, Table 4 shows the levels of the SEDI, government debt and net lending for the new Member States in 1999. In addition, at the bottom of the table, the average values of the respective variables are shown for the cohesion countries, as well as for the EU-11 Member States. The finding of very similar levels of government debt and net lending adds to the relevance of using the cohesion countries in the analysis.

		1999	
Country	SEDI	Government net lending	Government debt
Slovenia	0.558	-2.1 %	24.9 %
Hungary	0.549	-5.6 %	61.2 %
Czech Republic	0.494	-3.6 %	13.4 %
Latvia	0.487	-4.9 %	12.6 %
Poland	0.476	-1.4 %	40.3 %
Slovakia	0.450	-6.4 %	47.2 %
Lithuania	0.438	-5.6 %	23.0 %
Estonia	0.425	-3.7 %	6.0 %
Bulgaria	0.387	0.4 %	79.3 %
Romania	0.332	-4.5 %	24.0 %
Avg. NMS+AC	0.460	-3.7 %	33.2 %
Avg. Cohesion countries in 1980	0.376	-3.9 %	35.9 %
Avg. EU11 in 1980	0.454	-1.2 %	44.4 %

Table 4. Socio-Economic Development Index, government net lending and level of government debt. Sources: Authors' calculations and Eurostat.

5.2.2 Model Specifications

In this sub-section, the model specifications are discussed. Our methodology is similar to one used by Alesina *et al.* (2002), where the authors investigated the effects of fiscal policy on investment and profits of firms. They regressed profits on measures of government expenditure and revenues, and further used a breakdown of the series of government spending, similarly to our paper. In our analysis, we use cyclically adjusted variables in order to exclude the automatic response of fiscal variables to changes in economic conditions (such as the automatic stabilizer effects) and to measure the actual stance of fiscal policy. Some variables, such as the debt interest payments, are not, however, cyclically adjusted. Nevertheless, as they are to a large extent uncorrelated with business cycles, this point might be of minor importance. Finally, like Gali and Perotti (2003), we include the public debt to GDP ratio in our regressions, in addition to the government spending and revenue variables. We use ratios of the fiscal variables to potential GDP (and to trend GDP as a robustness test[17]) in the estimations.

Following the OECD data structure, we use a breakdown of government expenditure and revenues as follows. In the two most basic specifications, the independent variables are primary government balance and debt, and net lending and debt, respectively. Government net lending can be disaggregated into current receipts less current disbursements (excluding gross interest payments) less net capital outlays. Furthermore, the variable current receipts is disaggregated into taxes and received social security contributions. We use these fiscal variables with the public debt to GDP ratio to explain the evolution of the Socio-Economic Development Indicator of Greece, Ireland, Portugal and Spain for the period 1980-1999. In addition, the same estimation was conducted for the EU-11 Member States, as a control group, and as a robustness test for the entire sample of EU-15. In order to tackle the possible endogeneity issue, we used the instrumental variables estimation method (two stage least squares), using the first and second lags of the independent variables as instruments.[18] Linear and quadratic trends were also included in the models, as well as constant terms.

Even if changes in the variables in our index are to a large extent (or even exclusively) determined by fiscal policy, providing support for our model,

[17]The trend GDP was estimated for each country by regressing the log of real GDP against a constant, a linear and a quadratic trend.

[18]To increase the robustness of the results, we also estimated fixed effects models and obtained qualitatively very similar results. These results are available on request.

the time frame of the impacts on some variables could be questioned. For example, a reduction in the child mortality rate certainly reflects a longer term commitment in health care by the public sector than one captured by yearly changes in fiscal policy. However, even if the impact on some variables would only arise after a longer time period, our approach can be defended by the fact that countries have generally followed 'trends' in fiscal policies: years of fiscal profligacy are generally followed by years of fiscal consolidation.[19]

5.2.3 About the Data

For the development and fiscal indicators, the data are annual. The SEDI is constructed using data from the World Bank WDI 2003 database mentioned earlier.[20] The data for fiscal policy variables are obtained from the OECD Economic Outlook 75 database. Other data source is the Eurostat for the fiscal variables in Table 4. In the estimations, we used the STATA 8.2 statistical software.

The limited dimension of the panel of observations (4 countries × 20 annual observations for the fiscal series) creates problems for the evaluation of the time series properties of the series. On one hand, the number of time series observations is small to apply the single time series unit root tests, such as augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) tests. On the other hand, the number of cross-sections is also rather small to properly apply panel unit root tests.[21] However, using panel unit root tests can still be considered one way of increasing the power of the univariate tests, as stated by Maddala and Wu (1999). We chose to use the panel unit root tests by Levin, Lin and Chu (2002) (LLC test), given that also the cross-sectional dimension of the panel is limited.

The LLC test is based on an analysis of the following equation:

$$\Delta y_{i,t} = \alpha_i + \delta_i t + \theta_t + \rho_i y_{i,t-1} + \varsigma_{i,t}, \tag{5.3}$$

where $i = 1, 2, ..., N$ and $t = 1, 2, ..., T$. This model allows for fixed effects (α and θ) and unit-specific time trends. The unit-specific fixed effects are an important source of heterogeneity, since the coefficient of the lagged

[19]An example is Greece, where the budget deficit worsened from 1980 to 1990, then declined modestly until 1995, and fell at an accelerated pace until 1999.

[20]The data used is described in more detail in the Appendix.

[21]Therefore, we also made the panel unit root tests using the full sample of EU-15 countries.

dependent variable is restricted to be homogeneous across all units of the panel. The null hypothesis $H_0 : \rho_i = 0$ for *all* i is tested against the alternative $H_A : \rho_i = \rho < 0$ for *all* i (all series are stationary). Like most of the unit root tests in the literature, the LLC test assumes that the individual processes are cross-sectionally independent. Given this assumption, Levin *et al.* (2002) derived conditions under which the pooled OLS estimate of ρ will have a standard normal distribution under the null hypothesis.

When the LLC panel unit root test was applied to the fiscal series with respect to potential GDP, in 8 cases out of 9, the null hypothesis that *all four countries had a unit root* in their series at hand was rejected at a minimum of 5 percent level of significance against the alternative that *all* countries are stationary. Only in the case of government debt, the null hypothesis was not rejected at the conventional levels of significance. When the LLC test was applied to series with respect to trend GDP, the null hypothesis of a unit root could be rejected for all series at a minimum of 5 percent level of significance.[22] When the LLC tests were conducted for the entire sample of EU-15, the null hypothesis could be rejected for all series at the minimum of 5 percent level of significance, with series expressed both as shares to trend and to potential GDP. Finally, the SEDI variable was found to be stationary or trend stationary at the conventional levels of significance for all the country groups.

We can claim that these results would seem to justify estimating the models in levels instead of differences. Furthermore, there is a trade-off between differencing the series and losing information, and estimating the series in levels with a small possibility of (co)integrated series. In our case, it is hard to justify the usefulness of applying a panel cointegration analysis for this simple study with the limited panel of observations. Therefore, we proceeded with our analysis treating the variables as stationary or trend stationary. As a robustness check, we also estimated the models in first differences, but no conclusions could be drawn from those estimates. Finally, we should point out that our method is very similar to the one used in Alesina *et al.* (2002), where the authors estimated their models using the fiscal data from the same source, treating the variables as stationary.

[22] The null hypothesis for the variable "net capital outlays to trend GDP" was rejected at the 10 % level of significance.

5.2.4 Evolution of the Main Variables

Chart 1 below depicts the evolution of the Socio-Economic Development Index in the cohesion countries during 1980-1999. The SEDI variables of the EU-11 Member States, as well as of the new Member States are presented in the Appendix.

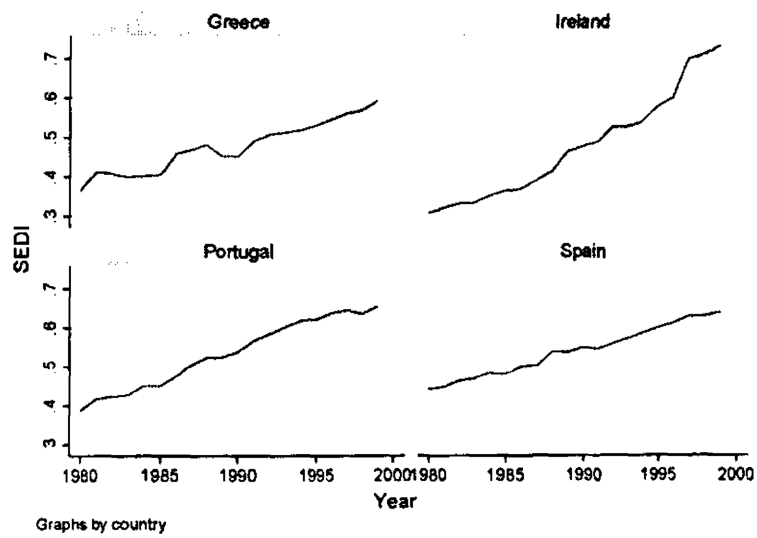


Chart 1. Socio-Economic Development Indicator for cohesion countries 1980-1999.

According to our indicator, Ireland was, in 1980, at the lowest level of socio-economic development of the cohesion countries. However, Ireland also had the fastest development rate: its socio-economic conditions improved during the sample period by a total of 138.3 percent, as measured by our index. Similarly, the smallest change in the development index, about 45.1 percent, took place in Spain that had the highest level of development in 1980. The lowest level of development in the EU-15 in 1999 was, according to our results, in Greece, where the index stood at 0.592. Finally, the average annual growth rate of the SEDI in the cohesion country group was 0.0147 SEDI units in 1980-1999.

During the sample period, the economies under study went through a notable fiscal consolidation. Charts 2 and 3 depict the evolution of government net lending relative to GDP, as well as the development of gross government debt to GDP in 1980-1999.

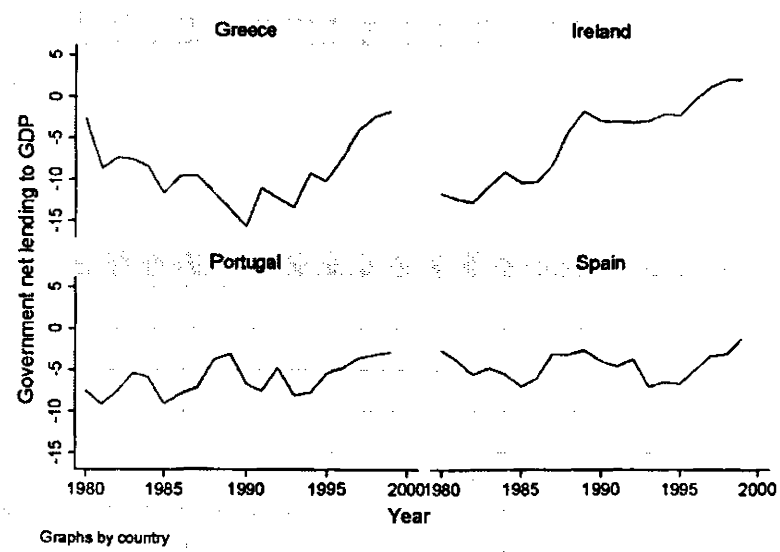


Chart 2. Government net lending to GDP for cohesion countries 1980-1999.

In Ireland, after peaking at -12.7 percent of GDP in 1982, government net lending significantly increased. Laski and Römisch (2003) reported that this was due to government expenditure growing more slowly than the GDP, and government revenues growing at a faster pace. Finally, government net lending turned positive in Ireland in 1997. In Greece, the budget deficit initially worsened rapidly and net lending reached -15.9 percent of GDP in 1990. There was a modest decline in the deficit until 1995, which was then followed by a faster improvement in the fiscal position, with net lending amounting to -1.8 percent of GDP in 1999. According to Laski and Römisch (2003), the average tax rate in Greece increased from 8.5 percent to 14.6 percent of GDP between 1995-2000. Similarly to the other economies under study, Portugal started from a very high budget deficit in the early 1980s (net lending in 1981 stood at -9.2 percent of GDP), whereas after that the budget deficit, expressed in terms of net lending, slowly declined to -2.9 percent of GDP in 1999. In Spain, the deficit in net lending rose first in the early 1980s, decreased somewhat in the late 1980s, and then rose rapidly to reach almost 7 percent of GDP in 1993. After that, fiscal consolidation was very fast and net lending amounted to -1.2 percent of GDP in 1999.

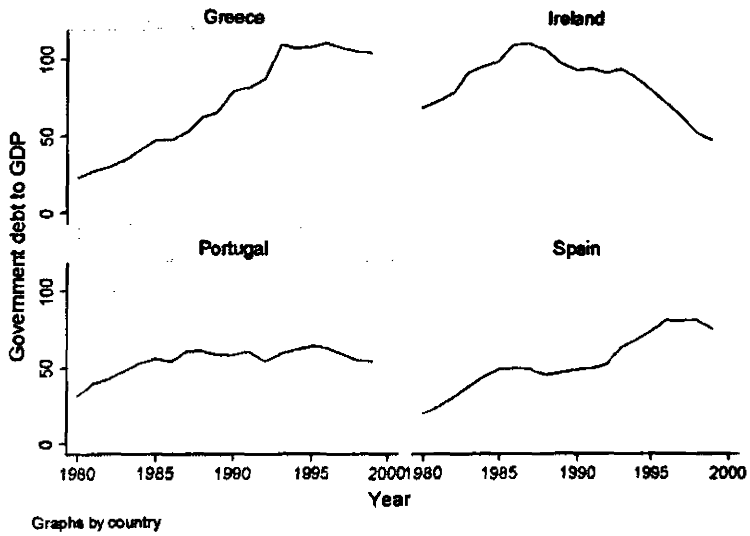


Chart 3. Government debt to GDP for cohesion countries 1980-1999.

Government debt to GDP was on a steadily increasing path during the examination period in Spain, Greece and to some extent also in Portugal. However, the deterioration in the debt to GDP ratio was notable in Greece in the early 1990s (from 66 to 110 percent of GDP between 1989 and 1993). In contrast, the debt to GDP ratio decreased from 96.5 to 49.3 percent of GDP during 1993-1999 in Ireland.

5.3 Empirical Results

5.3.1 Socio-Economic Development and Fiscal Policy

The regression results for the cohesion countries are presented in Table 5. The dependent variable is the natural logarithm of the Socio-Economic Development Indicator, while the independent variables are different fiscal measures in levels. The LOG-LIN models were estimated using fiscal variables with respect to potential GDP (calculated by the OECD). In the models, the level of government debt and its accumulation (with respect to different subcomponents) were controlled for.

TWO STAGE LEAST SQUARES LOG-LIN model for cohesion countries

Dependent variable: ln of SEDI, independent variables are used as instruments lagged 1 and 2 time periods

	1	2	3	4	5
Government Debt to Potential GDP	-0.4430***	-0.3660***	-0.3416***	-0.4070***	-0.3415***
	[0.0205]	[0.0188]	[0.0676]	[0.0331]	[0.0148]
Primary Government Balance to Potential GDP	1.3098***				
	[0.1885]				
Net Lending to Potential GDP		0.8067***			
		[0.1548]			
Current Receipts to Potential GDP			-0.1969	0.2476	
			[0.7901]	[0.3435]	
Total Expenditure to Potential GDP			-0.2871		
			[0.5669]		
Current Disbursements Excl. Interest Payments to Potential GDP				-0.6159**	-0.5930***
				[0.2625]	[0.1048]
Net Capital Outlays to Potential GDP				-1.8466***	-1.6444***
				[0.2867]	[0.2983]
Total Taxes to Potential GDP					-0.7053***
					[0.2154]
Social Contributions Received to Potential GDP					0.1597
					[0.2170]
Observations	72	72	72	72	72
R-squared	0.958	0.956	0.943	0.969	0.976
Hansen J-statistics	0.509	0.983	3.574	3.333	7.577
P-value	0.775	0.612	0.311	0.504	0.181

Huber/White robust standard errors in brackets

* significant at 10%; ** significant at 5%; *** significant at 1% level

Table 5. Estimation results for cohesion countries.

According to the results, an improvement in the ratio of primary government balance to GDP by 1 percentage point would increase the Socio-Economic Development Index by 1.31 percent *ceteris paribus*. Similarly, the coefficient for net lending is 0.81, and for current disbursements of a magnitude between -0.59 and -0.62. Furthermore, a decrease in the ratio of government debt to GDP by 1 percentage point would increase the Socio-Economic Development Index by 0.34 percent to 0.44 percent *ceteris paribus*. Both lower government spending and total taxes increase the SEDI, suggesting that socio-economic development benefits from the down-sizing of the public sector. A similar result is also suggested by the theoretical transition model by Chadha and Coricelli (1997). Contrasting evidence is found with respect to models where public capital investments contribute to aggregate production, as we find a significant negative coefficient on net capital outlays.[23]

Similar models were estimated for the group of the EU-11 Member States, and the results are reported in Table 6. Again, a reduction of government

[23] However, a possible explanation could be that if the level of government capital stock and investment are higher than the social optimum, then the tax burden on firms and citizens may also be too high and welfare could be improved by decreasing the government capital stock and spending.

debt improves our development index, but the coefficients are substantially lower for the other EU-11 Member States than for the cohesion countries. In addition, an increase in government net lending has a positive impact on the SEDI. However, the coefficient is only marginally significant at 10 percent level, while the coefficient for primary government balance is not statistically significant. These results implicate that fiscal consolidation would be more prominent in promoting socio-economic development in the cohesion countries. In contrast to our findings for the cohesion countries, an increase in government spending and taxes is found to be beneficial for socio-economic development in the other EU-11 Member States. This may be an indication of the public sector expenditure being more efficiently used in the EU-11 Member States. Then, fiscal consolidation would be beneficial for socio-economic development when conducted through an increase in revenues rather than through cuts in expenditure. As the literature emphasizes, the size and persistence of the fiscal adjustment, its composition and the initial state of public finances are important factors in determining the outcome of the economic policy. However, as the government debt levels for the cohesion countries were actually lower, but the net lending variables more strongly in deficit than the ones for the EU-11 Member States in the 1980s, it is likely to be the initial level of government net lending that causes the differing impacts of fiscal policy on socio-economic development between these two groups. Finally, it is important to note that we found the levels of government debt and net lending, as well as socio-economic development of the cohesion countries in the 1980s to have been strikingly close to those of the new Member States in 1999.

TWO STAGE LEAST SQUARES LOG-LIN model for EU-11 countries

Dependent variable: ln of SEDI. independent variables are used as instruments lagged 1 and 2 time periods

	1	2	3	4	5
Government Debt to Potential GDP	-0.0863***	-0.0484*	-0.0561***	-0.0595***	-0.0482***
	[0.0181]	[0.0261]	[0.0162]	[0.0169]	[0.0151]
Primary Government Balance to Potential GDP	-0.0067				
	[0.3502]				
Net Lending to Potential GDP		0.4881*			
		[0.2767]			
Current Receipts to Potential GDP			0.0227	-0.1025	
			[0.3091]	[0.3410]	
Total Expenditure to Potential GDP			0.8286**		
			[0.3391]		
Current Disbursements Excl. Interest Payments to Potential GDP				0.9124**	0.7204***
				[0.3568]	[0.1028]
Net Capital Outlays to Potential GDP				-0.7422	-0.1361
				[0.4627]	[0.4906]
Total Taxes to Potential GDP					1.0013***
					[0.2811]
Social Contributions Received to Potential GDP					0.1108
					[0.1062]
Observations	170	170	170	170	170
R-squared	0.721	0.731	0.833	0.834	0.852
Hansen J-statistics	1.194	3.413	0.384	0.918	5.133
P-value	0.550	0.181	0.944	0.922	0.400

Huber White robust standard errors in brackets

* significant at 10%; ** significant at 5%; *** significant at 1% level

Table 6. Estimation results for EU-11 countries.

To test for the robustness of the results, we also estimated the previous models for the whole sample of EU-15 Member States. The results from this estimation are reported in Table 7. This specification, even if it does not emphasize the differences between the cohesion countries and the EU-11 Member States, may be econometrically preferable due to a bigger sample size.[24] All in all, the results are in line with the ones previously reported. Similarly to the case of the cohesion countries and the EU-11 Member States, reductions in government debt increase socio-economic development. Furthermore, an increase in total expenditure now increases the development index, suggesting that results for the EU-11 Member States (excluding the cohesion countries) are dominating the findings from this specification. Notably, the effects of fiscal consolidation on socio-economic development are again weaker in terms of the estimated coefficients (with the exception of the net lending and net capital outlays variables) than in the case of the cohesion countries.

[24] However, in the second specification where government debt and net lending are used as independent variables, the Hansen J-test for exogeneity of our instruments is rejected.

TWO STAGE LEAST SQUARES LOG-LIN model for EU-15 countries
Dependent variable: ln of SEDI, independent variables are used as instruments lagged 1 and 2 time periods

	1	2	3	4	5
Government Debt to Potential GDP	-0.1549***	-0.0431*	-0.1161***	-0.1316***	-0.1380***
	[0.0254]	[0.0247]	[0.0202]	[0.0212]	[0.0201]
Primary Government Balance to Potential GDP	0.9430***				
	[0.3519]				
Net Lending to Potential GDP		1.2947***			
		[0.2134]			
Current Receipts to Potential GDP			-0.0942	-0.2537	
			[0.2828]	[0.2723]	
Total Expenditure to Potential GDP			0.9233***		
			[0.2906]		
Current Disbursements Excl. Interest Payments to Potential GDP				0.8653***	0.5166***
				[0.2788]	[0.0776]
Net Capital Outlays to Potential GDP				-1.8695***	-2.0860***
				[0.3858]	[0.4299]
Total Taxes to Potential GDP					0.2680
					[0.2529]
Social Contributions Received to Potential GDP					0.2881***
					[0.1098]
Observations	242	242	242	242	242
R-squared	0.67	0.703	0.819	0.832	0.835
Hansen J-statistics	3.516	8.523	1.367	4.073	5.926
P-value	0.172	0.014	0.713	0.396	0.313

Huber White robust standard errors in brackets
* significant at 10%; ** significant at 5%; *** significant at 1%

Table 7. Estimation results for EU-15 countries.

As a further robustness test, we estimated the models with variables expressed as ratios to trend GDP. For all the country groups (EU-15, EU-11 and the cohesion countries), the main results, including the sizes of the statistically significant coefficients, remained broadly unchanged. For models expressed both as ratio to potential and trend GDP, the results were not robust to a first difference transformation. These results are available on request.

5.3.2 Implications for the New Member States

In this section, we use our results to assess the implications for the new Member States. We first discuss the time it would take for the new Member States to reach the average and lowest welfare levels of the EU-15 in 1999, the year when the single currency was introduced. Then, we discuss some of the recent developments in the fiscal balances of the new Member States. It is important to note that this analysis and its implications should not be confused with the convergence criteria that are a prerequisite for euro area entry.

In Table 8, we list the number of years it would take for the new Member States to reach the development levels of the average EU-15 member and

Greece in 1999, assuming the new Member States developed at the average annual growth rate (0.0147 SEDI units) of the cohesion countries in 1980-1999,[25] and experienced similar paths of fiscal consolidation.

Country	Years to average EU-15 in 1999	Years to Greece in 1999
Slovenia	8.5	2.4
Hungary	9.1	3.0
Czech Republic	12.8	6.7
Latvia	13.4	7.2
Poland	14.1	8.0
Slovakia	15.8	9.7
Lithuania	16.7	10.5
Estonia	17.6	11.4
Bulgaria	20.1	13.9
Romania	23.9	17.7

Table 8. Socio-Economic Development convergence time.

It is clear from Table 8 that the convergence times vary significantly, depending on the level of development that is aspired to. In 1999, Slovenia was lagging behind the EU-15 average level of development by 8.5 years, but only 2.4 years behind the level in Greece. For Romania, the numbers of years are 23.9 and 17.7, respectively. We find that the convergence times in terms of socio-economic development are slightly lower than the often-investigated income convergence times. As an example, Fischer *et al.* (1998) examined how long it would take the transition countries of Eastern and Central Europe to close the income gap to the current EU countries, and arrived at an average time of 30 years. Similarly, Wagner and Hlouskova (2002) suggested that except for Slovenia and the Czech Republic, the average time it would take for the new Member States to achieve 70 percent or 80 percent of the enlarged EU's average GDP level is 30-40 years. However, one should note that the convergence times reported in Table 8 are the times required to reach the desired socio-economic development level that the EU-15 Member States had in 1999, not the catching-up times.

What do the predictions from the theoretical model by Chadha and Coricelli (1997) imply, if they are considered together with our results? First, as Coricelli (1998) has argued, the new Member States may experience some slowdown in their convergence process. This would be a response to fiscal constraints that have become tighter as convergence has progressed. If restructuring is still sought at a rapid pace, this may make it more difficult to keep

[25] We justify this assumption again by pointing to the similar development levels of the cohesion countries at the time of their EU accession and the new Member States in 1999.

fiscal balances in order. This would, according to our results, be detrimental in terms of welfare, as it would have a negative impact on the socio-economic development indicator. Second, unemployment is a major problem in most new Member States. Depending on how far the convergence process has progressed, there are differences in its impact on socio-economic development. If the convergence process is still at the initial stages, a fast restructuring would imply further unemployment, worsening fiscal balances and an adverse impact on the socio-economic development. However, it is more likely that in many new Member States the convergence has already progressed somewhat further. Then, advancements in the convergence process would instead increase output growth, decrease unemployment, improve the fiscal balances and, according to our results, have a positive impact on welfare.

In many new Member States, government spending and the general government deficit have recently increased significantly, especially in Hungary, Slovakia and the Czech Republic. Moreover, with high GDP growth lowering the value of the fiscal variables expressed as ratios to GDP, it becomes clear that the deficits have their origins in strong expenditure pressures and, as the EBRD (2003) points out, the deficits in these countries are largely structural in nature. Therefore, the task of reducing budget deficits and government debt levels will be a challenging one.

5.4 Conclusions

The aim of our paper was to examine the link between socio-economic development and fiscal policy. In order to achieve our aim, we first constructed a Socio-Economic Development Index (SEDI) and then regressed it on a number of fiscal variables, including variables from both the expenditure and revenue side of the government balance sheet, from Greece, Ireland, Portugal and Spain for 1980-1999. During this time period, the countries entered the European Union and started the necessary adjustments toward the single currency, introduced in 1999. We then used the results from our instrumental variables regressions to evaluate the implications for the new Member States. Finally, we also calculated how long it would take for the new Member States to achieve the EU benchmark levels in 1999 in terms of the development indicator, assuming a speed of development of the above countries during 1980-1999. The times varied from 8.5 years (Slovenia) to 24 years (Romania).

Our results show that fiscal consolidation would be beneficial for socio-economic development in the medium term. In line with the literature on the effects of fiscal consolidation on economic output and growth, we find that fiscal retrenchment through a lower government debt and an improved net lending position would be advantageous to socio-economic development. The effects of fiscal consolidation in promoting socio-economic development are found to be much stronger for the cohesion countries than for the other EU-15 Member States. Whereas an overall down-sizing of the public sector was found to improve socio-economic development in the cohesion countries, in the other EU-15 Member States increases in government current disbursements were found to have beneficial effects on development, suggesting a more efficient public sector in those economies. All in all, the results could be seen to support maintaining the Stability and Growth Pact or an equivalent intergovernmental device to curb public spending and debt.

Bibliography

[1] Afonso, A., Schuknecht, L., Tanzi, V., 2003. Public Sector Efficiency: An International Comparison. ECB Working Paper no. 242.

[2] Alesina, A., Ardagna, S., 1998. Tales of Fiscal Adjustment. Economic Policy 27, 487-546.

[3] Alesina, A., Ardagna, S., Perotti, R., Schiantarelli, F., 2002. Fiscal Policy, Profits, and Investment. American Economic Review 92 (3), 571-589.

[4] Alesina, A., Perotti, R., 1995. Fiscal Expansions and Adjustments in OECD Countries. Economic Policy 21, 205-240.

[5] Budina, N., van Wijnbergen, S., 1997. Fiscal Policies in Eastern Europe. Oxford Review of Economic Policy 13, 47-64.

[6] Chadha, B., Coricelli, F., 1997. Fiscal Constraints and the Speed of Transition. Journal of Development Economics 52, 221-249.

[7] Coricelli, F., 1998. Macroeconomic Policies and the Development of Markets in Transition Economics. Central European University Press, Budapest.

[8] Dewatripont, M., Roland, G., 1992. The Virtues of Gradualism and Legitimacy in the Transition to a Market Economy. The Economic Journal 102, 291-300.

[9] Diamond, P., 1965. National Debt in a Neoclassical Growth Model. American Economic Review 55 (5), 1126-1150.

[10] EBRD, 2003. Transition Report Update. EBRD Publications.

[11] European Commission, 2003. Public Finances in EMU 2003. Directorate-General for Economic and Financial Affairs. May 2003.

[12] Fischer, S., Sahay, R., Vegh, C., 1998. How Far is Eastern Europe from Brussels? IMF Working Paper no. 98/53.

[13] Gali, J., Perotti, R., 2003. Fiscal Policy and Monetary Integration in Europe. Economic Policy 18 (37), 533-572.

[14] Giavazzi, F., Pagano, M., 1990. Can Severe Fiscal Adjustments Be Expansionary? Tales of Two Small European Countries. NBER Macroeconomic Annual 5.

[15] Giavazzi, F., Pagano, M., 1996. Non-Keynesian Effects of Fiscal Policy Changes: International Evidence and Swedish Evidence. Swedish Economic Policy Review 3 (3), 67-103.

[16] Giavazzi, F., Jappelli, T., Pagano, M., 2000. Searching for Non-Linear Effects of Fiscal Policy: Evidence from Industrial and Developing Countries. European Economic Review 44 (7), 1259-1289.

[17] Laski, K., Römisch, R., 2003. From Accession to Cohesion: Ireland, Greece, Portugal and Spain and Lessons for the Next Accession. WIIW Research Report no. 298, July 2003.

[18] Levin, A., Lin, C. F., Chu, C. S., 2002. Unit Root Tests in Panel Data: Asymptotic and Finite-Sample Properties. Journal of Econometrics 108, 1-24.

[19] Maddala, G. S., Wu, S., 1999. A Comparative Study of Unit Root Tests with Panel Data and a New Simple Test. Oxford Bulletin of Economics and Statistics, Special Issue, 631-652.

[20] Modigliani, F., 1961. Long-Run Implications of Alternative Fiscal Policies and the Burden of the National Debt. Economic Journal 71 (284), 730-755.

[21] Pelkmans, J., Gros, D., Nunez Ferrer, J., 2000. Long-Run Economic Aspects of the European Union's Eastern Enlargement. Netherlands Scientific Council for Government Policy Working Paper no. 109.

[22] Perotti, R., 1999. Fiscal Policy in Good Times and Bad. Quarterly Journal of Economics 114, 1399-1436.

[23] Perotti, R., 2002. Estimating the Effects of Fiscal Policy in OECD Countries. ECB Working Paper no. 168.

[24] Pirttilä, J., 2001. Fiscal Policy and Structural Reforms in Transition Economies - an Empirical Analysis. Economics of Transition 9 (1), 29-52.

[25] Rzonca, A., Cizkowicz, P., 2003. A Comment on The Relationship Between Policies and Growth in Transition Countries. Economics of Transition 11 (4), 743-748.

[26] Saint-Paul, G., 1992. Fiscal Policy in an Endogenous Growth Model. Quarterly Journal of Economics 107, 1243-1259.

[27] United Nations, 2001. Human Development Report 2001. Oxford University Press, Oxford.

[28] United Nations, 2003. Human Development Report 2003. Oxford University Press, Oxford.

[29] Vaitilingam, R. (Ed.), 2002. Who's Afraid of the Big Enlargement? Economic and Social Implications of the European Union's Prospective Eastern Expansion. CEPR Policy Paper no 7.

[30] Wagner, M., Hlouskova, J., 2002. The CEEC10s Real Convergence Prospects. CEPR Discussion Paper no. 3318.

5.5 Appendices

The Appendix describes the data sources and other data related issues.

First, all the fiscal variables were obtained from the OECD Economic Outlook 75 database. In the baseline models, the fiscal variables were transformed into ratios to potential GDP (as calculated by the OECD). As a robustness test, the models were estimated also using fiscal variables as ratios to real trend GDP that was obtained by regressing the real GDP (deflated by the GDP deflator, obtained from the OECD) against a constant, a linear and a quadratic trend.

Second, the variables for the Socio-Economic Development Indicator (SEDI) were obtained from the World Bank World Development Indicators (WDI) 2003 database. The SEDI was constructed using data for the EU-15 countries (1992-1999 for Germany, 1980-1999 for the other EU-15 Member States) and for the new Member States (1992-1999). Due to limited data availability, some variables included in the SEDI needed to be interpolated or extrapolated, as follows (missing years in parentheses):

The variable "air passengers carried" was missing for Slovakia (1992).

The variable "rail passenger kilometers" was missing for Denmark (1999), Ireland (1999), and the Netherlands (1999).

The variable "infant mortality rate" was missing for Belgium (1996, 1999), Finland (1984, 1988), Greece (1999), and Italy (1999).

The variable "immunisation DPT" was missing for Austria (1980), Belgium (1980), the Czech Republic (1992), Denmark (1980), Estonia (1992), Finland (1980), France (1980), Italy (1980-83), Slovakia (1992-93), Spain (1980-83), Sweden (1980), and the United Kingdom (1980).

Finally, the variables "primary and tertiary school enrollment" were available before 1990 for all countries only in 1980, 1985 and 1990. Therefore, the variables were linearly interpolated for all countries for 1981-1989. In addition, "primary school enrollment" was also missing in Belgium (1997), Ireland (1998), Poland (1998), and "tertiary school enrollment" in Belgium (1998), Finland (1998-99), Germany (1999), Greece (1998-99), and Slovenia (1999).

Note, however, that the missing observations for the new Member States had no impact on the actual panel estimations, as the models were estimated using data for the cohesion countries and the other EU-15 Member States only. Finally, the natural logarithm of SEDI was used in the models.

Other data source is the Eurostat for the fiscal variables in Table 4.

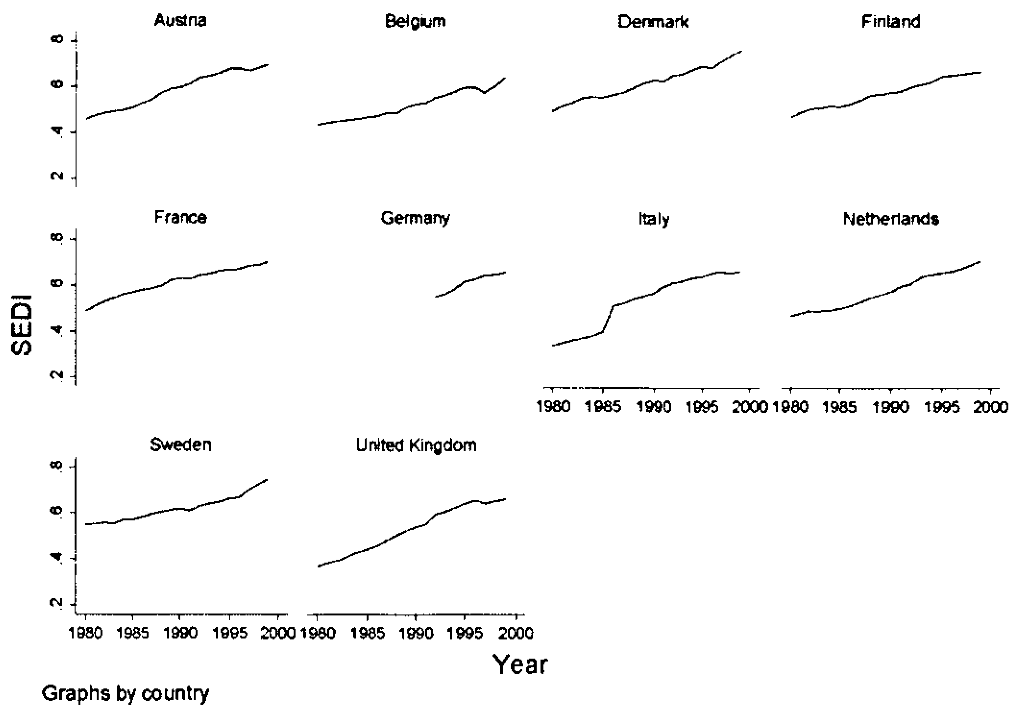


Figure: Socio-economic development indicator for EU-11 countries 1980-1999.

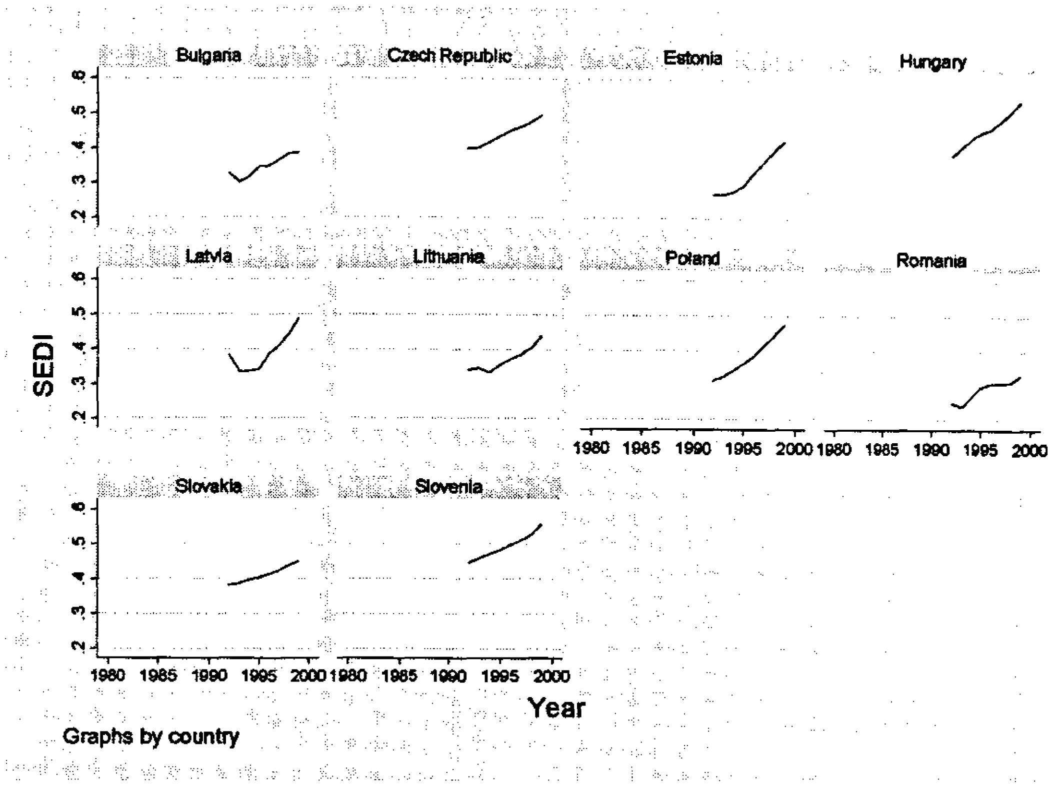


Figure: Socio-economic development indicator for new Member States 1980-1999.

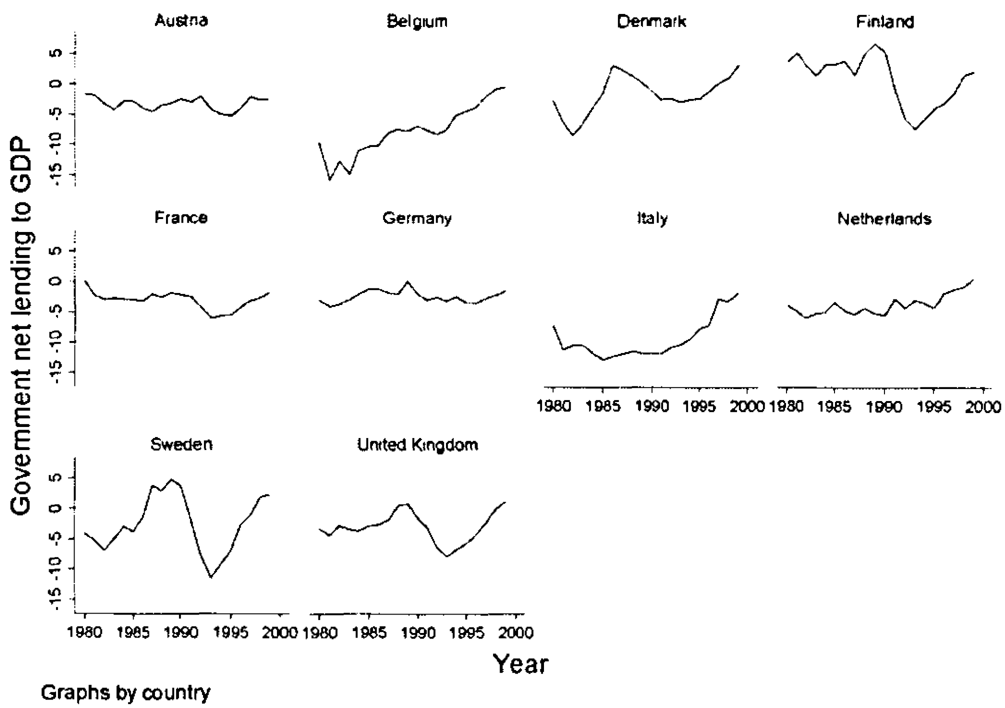


Figure: Government net lending for EU-11 countries 1980-1999.

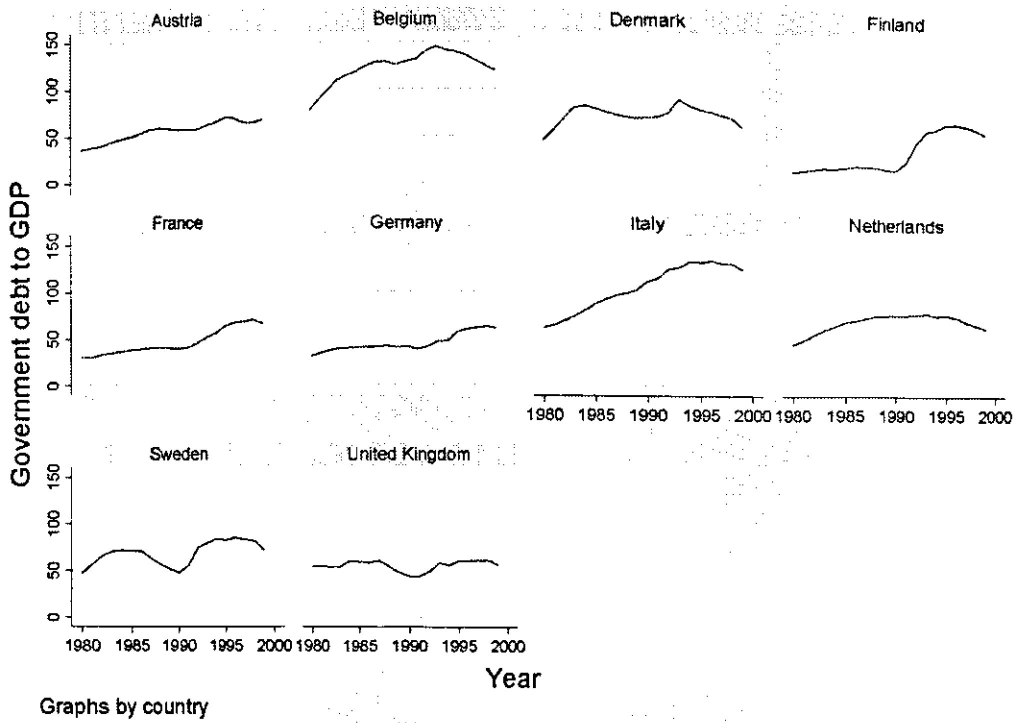


Figure: Government debt to GDP for EU-11 countries 1980-1999.