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Monetary Shocks and Stock Returns: Identification Through the Impossible Trinity

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Abstract:

This paper attempts to identify how monetary policy shocks affect stock prices by using Mundell and Fleming's theory of the "Impossible Trinity." According to this theory, it is impossible to simultaneously have a fixed exchange rate, free capital movement (an absence of capital controls), and an independent monetary policy. The authors present evidence that Hong Kong's monetary policy is heavily dependent on the monetary policy of the United States, a stance which is consistent with this theory because the HK dollar has been pegged to the U.S. dollar since 1983 and Hong Kong does not impose any capital controls. As a result, the Federal Reserve's monetary policy actions can be considered as exogeneous shocks to the Hong Kong economy. Recognizing this relationship helps us solve the endogeneity problem inherent in the studies examining the relationship between stock prices and monetary policy shocks. This is the first paper that presents evidence of severe omitted variable bias in the event studies focusing on the relationship between monetary policy and stock returns. The authors also suggest a way to remedy this bias.

JEL Classifications: E4 (E44), E5 (E52, E58), G1 (G12, G15, G18)

Keywords: stock prices; monetary policy; simultaneity, omitted variables

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1 Introduction

What is the effect of monetary policy on stock prices? The answer to this question is important for both investors and policymakers. For investors, it is important to know the extent to which their stock market holdings are exposed to monetary policy shocks. For policymakers, it is crucial to understand how monetary policy affects the real economy through its influence on stock prices.

As illustrated in Rigobon and Sack (2004), there are two major identification difficulties in the literature that studies the relationship between stock prices and monetary policy. The endogeneity (simultaneity) problem arises from the joint determination of monetary policy and stock returns because monetary policy can simultaneously react to changes in stock prices.[1] The omitted variable problem arises from the possibility that stock returns and monetary policy variables may be jointly reacting to some other macroeconomic variables that would cause a bias even in the absense of the endogeneity problem.[2]

We solve the endogeneity problem by using the Impossible Trinity theory developed in Fleming (1962) and Mundell (1963). According to the Impossible Trinity theory, it is impossible to simultaneously have a fixed exchange rate, free capital movement (an absence of capital controls), and an independent monetary policy. Hong Kong offers a clear example of Mundell and Fleming's Impossible Trinity theory. First, Hong Kong imposes no restrictions on capital flows or on the trading of financial assets. Second, as shown in Figure 1, the Hong Kong Monetary Authority (HKMA) has successfully implemented a fixed exchange rate for the HK dollar/U.S. dollar since October 1983.[3] Since the establishment of the Linked Exchange Rate System, the Hong Kong dollar exchange rate has remained stable in the face of various shocks. It was unaffected by the

[1] See Rigobon and Sack (2003) for evidence of the endogeneity problem.

[2] Using data published by *Money Market Services*, we find that at least 78 out of 177 monetary policy announcement dates between 1989 and 2008 overlap with other macroeconomic announcements that may influence both stock prices and monetary policy.

[3] The HKMA guaranteed to exchange U.S. dollar into HK dollar, or vice versa, at a predetermined rate until 2005. Since May 18, 2005, the HKMA has set a very narrow band of 7.85 as an upper limit and 7.75 as a lower limit for the HK dollar.

1987 stock market crash, the 1990 Gulf War, the 1992 turmoil over the European Exchange Rate Mechanism, the 1994–1995 Mexican currency crisis, the 1997–1998 Asian financial crisis, the 9/11 incident, and the 2008 global financial crisis.

Since Hong Kong has free capital movement and a fixed exchange rate, the Impossible Trinity theory suggests that Hong Kong's monetary policy depends on U.S. monetary policy. Figure 2 provides evidence of the close monetary policy relationship between Hong Kong and the United States — movements in the Hong Kong base rate closely follow movements in the federal funds target rate.[4] Given that changes in the Hong Kong base rate closely follow changes in the fed funds target rate, and that the U.S. Federal Open Market Committee (FOMC) does not base its monetary policy on the stock price movements in Hong Kong, we can conclude that unexpected changes in the federal funds target rate are exogenous shocks for Hong Kong's monetary policy.[5]

To be sure, there have been episodes, particularly during the Asian financial crisis, where the markets have tested the HKMA's ability to stick to a fixed exchange rate. However, the HKMA has managed to successfully survive these episodes. Moreover, the HKMA's policies during the Asian crisis were mostly independent of the FOMC decisions except when the HKMA adjusted its base rate to reflect the changes in the federal funds target rate.[6] Therefore, the federal funds target rate changes during this period still can be considered as exogenous monetary policy shocks to the Hong Kong economy. Nevertheless, we control for these time periods in our robustness checks to

[4]Readers may be aware of the fact that the base rate is set mechanically by the HKMA through a transparent formula, which is the U.S. federal funds target rate plus some predetermined premium. This premium is designed to discourage the banks from accessing the discount window and is not significant for our analysis, since our empirical analysis focuses on the *changes* of the interest rate imposed by monetary policy rather than *levels*.

[5]The Impossible Trinity also can be applied in the context of other countries that have an exchange rate peg. We focus on Hong Kong because it has both a clearly defined fixed exchange rate and a well-developed stock market. Other candidates seem to miss one of these qualities. For example, Singapore maintains a currency peg, but rather than being a hard peg it is an adjustable peg in the form of a monitoring band arrangement with the central parity based on an undisclosed trade-weighted currency basket. Another example is Bermuda, where the Bermuda dollar is at par with the US dollar, but the stock market is not well-developed.

[6]Using daily data from the HKMA, we find that the HKMA always matched the FOMC target rate changes immediately, with only one exception. On September 28, 1998, the FOMC decreased the target rate by 25 basis points but the HKMA decreased the base rate by 12.5 basis points on impact. However, within three business days the HKMA followed this move with another 12.5 basis point decrease, thereby matching the full 25 basis point change in the fed funds target rate.

confirm the stability of our results.

So far, we have focused on the endogeneity problem explicitly. But what about the omitted variable problem? Of course, the Impossible Trinity theory does not provide the ultimate solution to the omitted variable problem because there may be global shocks directly affecting the U.S. and Hong Kong stock markets, in addition to their indirect effects acting through U.S. monetary policy. We use two steps to address and solve the omitted variable problem. First, we show that a simple regression of Hong Kong stock price growth on monetary policy surprises can suffer severely from omitted variable bias. Second, we present evidence that this bias disappears once we add U.S. stock returns as an additional control variable in the regression.[7] To the best of our knowledge, this is the first paper to provide explicit evidence of omitted variable bias in regressions where a measure of stock returns is a dependent variable and to offer a method that directly addresses the problem.[8]

Our paper contributes to the literature that studies the relationship between monetary policy and stock returns. Rigobon and Sack (2004) develop an estimator that identifies the response of asset prices based on the heteroskedasticity of monetary policy shocks on the event and pre-event dates. Gürkaynak, Sack, and Swanson (2005) use intraday data in a relatively narrow "event window" surrounding the FOMC's scheduled policy announcements, thereby distinguishing the impact of the policy change from the effect of news arriving earlier or later in the day. Bjørnland and Leitemo (2009) use both short-run and long-run restrictions in a VAR framework to control for endogeneity.[9]

[7]By addressing the omitted variable issue explicitly, we allow for the possibility that U.S. monetary policy responds to macroeconomic events that directly affect both the U.S. economy and the Hong Kong economy directly. Therefore, our identification mechanism only excludes the possibility that the Federal Reserve responds to idiosyncratic movements in Hong Kong stock prices, which is a realistic depiction of FOMC decisions.

[8]Rigobon and Sack (2004) find a significant bias for the response of Treasury yields to U.S. monetary policy shocks when one uses a standard OLS specification but they do not find a similar statistically significant bias for the response of stock prices.

[9]These papers follow a long line of preceding studies that look at the relationship between monetary policy and asset prices and address the identification problems to various degrees, see Geske and Roll (1983), Kaul (1987), Bomfim (2003), Bomfim and Reinhart (2000), Cochrane and Piazzesi (2002), Kuttner (2001), Bernanke and Kuttner (2005), Thorbecke (1997), Lee (1992), Patelis(1997), Fuhrer and Tootell (2003). See also Sellin (2001) for an earlier

Our empirical analysis is closely related to Bernanke and Kuttner (2005), who studied the reaction of the U.S. stock market to changes in the federal funds target rate. Following their method, we use changes in the federal funds futures' price on the dates of monetary policy announcements in order to identify surprise changes in the federal funds target rate. We follow this method because federal funds futures outperform target rate forecasts based on other financial market instruments or based on alternative methods, such as sophisticated time series specifications and monetary policy rules.[10] However, we use Hong Kong stock returns on these event dates, rather than U.S. stock returns, as the dependent variable and use U.S. stock returns to control for omitted variables. Therefore, our regressions do not suffer from the identification problems discussed in Rigobon and Sack (2004). Moreover, unlike previous studies, we present direct evidence for the omitted variable bias and provide evidence that our identification method explicitly addresses this bias.

Our method also has other advantages in comparison to previous studies. Unlike Rigobon and Sack (2004), our identification method does not assume that nonmonetary shocks and variables are homoscedastic.[11] Moreover, our identification mechanism allows us to omit some potential pitfalls introduced by high frequency intraday data. First, the higher volatility of high frequency data can cause an amplified errors-in-variables problem. Second, the high frequency stock price drift prior to FOMC announcements, documented by Lucca and Moench (2012), suggests that a very narrow event window can be misleading. Third, stocks do not seem to respond to FOMC announcements at the same speed. In particular, stocks with smaller market capitalizations have delayed reactions to monetary policy shocks, an effect which limits the ability of intraday data to capture the full effect of monetary policy shocks. This delayed response is particularly important for papers

survey.

[10]See Evans (1998) and Gürkaynak, Sack and Swanson (2007). Another advantage of looking at one-day changes in near-dated federal funds futures is that federal funds futures do not exhibit predictable time-varying risk premia (and forecast errors) over daily frequencies. See, for example, Piazzesi and Swanson (2008).

[11]Since U.S. monetary policy announcement dates between 1989 and 2008 overlap with other macroeconomic announcements at least 78 out of 177 times, the nonmonetary news may not be homoscedastic at event and pre-event dates, in contrast with the assumption of Rigobon and Sack (2004).

that use the event study method for cross-sectional comparisons of different firms' reactions to monetary policy. Fourth, even if the stock market and federal funds futures market immediately reacted to FOMC-related news, there can still be problems with identification if other important macroeconomic news was released earlier in the day and stocks did not react immediately.

The studies that employ high frequency data assume that a narrow window—for example, 15 minutes before and 45 minutes after the announcement—is enough to capture the full effect of the FOMC announcement because markets react to FOMC announcements very fast. Suppose that another macroeconomic announcement occurs earlier in the day, such as unemployment or Consumer Price Index (CPI) reports, news that markets do not immediately react to. Even if the markets react to an FOMC announcement immediately, a delayed response to other macroeconomic news implies that movements in S&P500 during the (-15m,$+45$m) event window are due not only to the FOMC announcement but are also partially due to a delayed response to this other macroeconomic news. This delayed response brings the identification problems back into the picture: To the extent that the macroeconomic news has been incorporated in the FOMC decision but not yet accounted for in stock prices, the endogeneity problem is back in focus because the response of stocks during the (-15m,$+45$m) event window is a combination of the responses to the FOMC announcement and the other macroeconomic news—although the FOMC announcement already incorporated the macroeconomic news and the associated delayed response of stocks. Also, to the extent that both the federal funds futures and stock markets continue to respond to the macroeconomic news earlier in the day during the (-15m,$+45$m) event window, the omitted variable problem is back in the picture because part of the correlation between stock prices and the fed funds futures price during the event window is attributable to their joint response to macroeconomic news. It is not easy to verify if these effects are unimportant.

2 Econometric Models: The Identification Problem Revisited

Hereafter, we omit time subscripts and constant terms in the econometric models for the sake of brevity. To be more precise, one can think of the variables as their de-meaned versions, given by the actual value minus the average value of each variable.

2.1 Endogeneity (Simultaneity) Problem

Monetary policy might respond to stock returns at the same time that the stock returns respond to monetary policy. Suppose Δs is the change in the U.S. stock price and Δi is the change in the federal funds target rate. Then we have the standard simultaneous equations problem,

$$\begin{aligned} \Delta i &= \beta \Delta s + \varepsilon \\ \Delta s &= \alpha \Delta i + \eta . \end{aligned}$$

If we use Ordinary Least Square (OLS) to estimate α in the second equation, we get

$$\operatorname{plim} \hat{\alpha}_{OLS} = \frac{cov(\Delta s, \Delta i)}{var(\Delta i)} = \alpha + \frac{cov\,(\eta, \Delta i)}{var\,(\Delta i)} \neq \alpha .$$

To find the magnitude of the bias, we first solve the above system for Δi,

$$\Delta i = \frac{\beta \eta + \varepsilon}{1 - \alpha \beta}.$$

The bias is then given by

$$\frac{cov\,(\eta, \Delta i)}{var\,(\Delta i)} = (1 - \alpha\beta)\,\frac{\beta \sigma_\eta}{\beta^2 \sigma_\eta + \sigma_\varepsilon},$$

where σ_x is the variance of variable x.

Our use of Hong Kong stock prices solves this problem because they do not enter into U.S.

monetary policy decisions directly, and the changes in U.S. monetary policy can be considered as exogenous shocks to the Hong Kong economy according to the Mundell-Fleming model. That is, if we let Δy be the change in stock prices in Hong Kong, we have

$$\Delta y = a\Delta i + w,$$

with $cov\,(\Delta i, w) = 0$. The estimation of this model via OLS gives

$$\text{plim}\,\hat{a}_{OLS} = \frac{cov(\Delta y, \Delta i)}{var(\Delta i)} = a + \frac{cov(w, \Delta i)}{var(\Delta i)} = a,$$

which is an unbiased estimate.

Although Hong Kong stock prices do not directly enter into FOMC policy decisions, they might be indirectly correlated with these decisions when there are global shocks affecting both variables. This is what we focus on next.

2.2 Omitted Variable Problem

Some economic news that affects monetary policy might also affect stock prices directly, in addition to the indirect effect of this news acting through monetary policy. This indirect effect can generate an omitted variable bias. To see this more clearly, suppose that the true econometric model is given by

$$\begin{aligned} \Delta i &= \gamma z + \varepsilon \\ \Delta s &= \alpha \Delta i + z + \eta, \end{aligned}$$

where z includes variables that affect stock prices directly, as captured by the second term, and indirectly through monetary policy. In this case, the OLS regression of Δs on Δi gives

$$\text{plim}\,\hat{\alpha}_{OLS} = \frac{cov(\Delta s, \Delta i)}{var(\Delta i)} = \alpha + \frac{cov\,(z, \Delta i)}{var\,(\Delta i)} = \alpha + \frac{\gamma \sigma_z}{\gamma^2 \sigma_z + \sigma_\varepsilon} \neq \alpha,$$

which is a biased estimate unless $\gamma = 0$.

This problem is not directly addressed by using Hong Kong stock market data. In particular, if there are any variables that affect both U.S. monetary policy and Hong Kong stock returns directly, that is, if

$$\Delta y = a\Delta i + ez + w$$

is the true model, and $e\gamma \neq 0$, a regression that does not include these variables, z, would still produce a biased estimate for a. Therefore, omitted variables can still pose a problem for our regressions, although the problem is likely less severe for Hong Kong stocks than for U.S. stocks because $e\gamma = 0$ is a weaker condition than $\gamma = 0$. In our analysis, we show that the omitted variable bias is potentially a very severe problem. We also show that using U.S. stock returns as an additional regressor can mitigate this problem.

3 Data

The data of our empirical study fall into two categories: indices of the U.S. and Hong Kong equity markets, and variables that represent changes in U.S. monetary policy.

As in Bernanke and Kuttner (2005), we use the total return on the CRSP value-weighted index as a measure of U.S. equity returns.[12] Our major indicator for the stock market performance in Hong Kong is the daily Hang Seng index (HSI).[13]

[12]The CRSP value-weighted index can be accessed through the CRSPSift system. Its INDNO is 1000200. See http://www.crsp.com/documentation/product/stkind/data_descriptions_guide.pdf for detailed information on index description and the calculation method for the index return.

[13]The Hang Seng index can be accessed through Bloomberg; the ticker is HSI Index. We use the growth of the HSI

One problem associated with estimating the market's reaction to monetary policy changes is that the equity market is not likely to respond to anticipated policy actions. To ease the problem, Bernanke and Kuttner (2005) adopt a method, proposed by Kuttner (2001), that separates the unexpected, or "surprise," component from the anticipated component of a monetary policy change—specifically, a change in the federal funds target rate. The identification of the surprise element in the target rate change relies on the price of 30-day federal funds futures contracts, a price which encompasses market expectations of the effective federal funds rate.

Following Bernanke and Kuttner's analysis, we define an event as either an FOMC meeting or an announced change in the funds target rate. Kuttner (2001) and Bernanke and Kuttner (2005) obtain the corresponding surprise change in the target rate by first calculating the change in the rate implied by the corresponding futures contract, given by 100 minus the futures contract price, and then scaling this result by a factor associated with the number of days of the month in which the event occurred. Accordingly, the unanticipated target rate change, for an event taking place on day d of month m, is given by

$$\Delta i^u = \frac{D}{D-d}(f^0_{m,d} - f^0_{m,d-1}),$$

where $f^0_{m,d} - f^0_{m,d-1}$ is the change in the current-month implied futures rate, and D is the number of days in the month. To suppress the end-of-month noise in the federal funds rate, the unscaled change in the implied futures rate is used as the measure of target rate surprise when the event occurs during the last three days of the month. If the event happens on the first day of the month, $f^1_{m-1,D}$ is used instead of $f^0_{m,d-1}$. The expected federal funds rate change is defined as the difference

index in U.S. dollar in our regressions to make the results comparable to studies that focus on the U.S. stock market. Using the HSI index in HK dollar instead has only a tiny quantitative effect on our results because the exchange rate fluctuations are negligibly small, which confirms that any minor movement in the secondary market exchange rate is not significantly correlated with U.S. monetary policy surprises. The Hong Kong stock market is closed at the time of scheduled FOMC announcements, since Hong Kong local time is twelve hours ahead of U.S. eastern daylight savings time. We adjust forward one day for the Hong Kong data.

between the actual change minus the surprise:

$$\Delta i^e = \Delta i - \Delta i^u,$$

where Δi is the actual federal funds rate change.

The data for the decomposition of the federal funds target rate changes can be obtained from Kenneth Kuttner's webpage.[14] Kuttner's dataset contains futures-based funds rate surprises on event days from June 1989 to June 2008, after which the Federal Reserve switched from announcing a specific target rate to announcing a range for the target rate. In our initial analysis of stock prices, there are three reasons why we focus primarily on the period between February 1994 to May 2005. First, starting in February 1994, the FOMC's policy of announcing target rate changes at pre-scheduled dates virtually eliminated the timing ambiguity associated with rate changes prior to this time period.[15] Second, Hong Kong switched to a narrow floating band policy on May 18, 2005. Third, during this time period, the Federal Reserve had the same Chairman, Alan Greenspan, and this continuity decreases possible contamination of our results due to a change in the policy regime.[16] We also check the robustness of our results by extending the dataset to include the period from June 1989 to June 2008.

3.1 The HIBOR vs. the Federal Funds Rate

Before our analysis of the stock index, we want to provide further evidence regarding the close relationship between U.S. monetary policy and overnight interest rates in Hong Kong. Figure 3 presents the overnight Hong Kong Interbank Offered Rate (HIBOR), the closest interest rate to the federal funds effective rate for Hong Kong. Although the HIBOR closely follows the federal funds target rate, it tracks neither the federal funds target rate nor federal funds effective rate. This

[14] http://econ.williams.edu/people/knk1/research

[15] Rigobon and Sack (2004) focus on the post-1994 period for the same reason.

[16] Alan Greenspan was succeeded by Ben Bernanke on February 1, 2006.

was particularly evident during the Asian financial crisis of late 1990s and after September 2003. This observation is confirmed by the comparison of adjusted R-squares in the left panel of Table 1. This pattern is to be expected because Hong Kong banks do not have direct access to the Federal Reserve facilities, as do U.S. banks.[17]

Nevertheless, for our identification mechanism to hold, that is, for federal funds target rate changes to be considered as exogenous monetary policy shocks to the Hong Kong economy, it is sufficient that a surprise change in the federal funds target rate causes a proportionate change in the HIBOR rate. The right panel of Table 1 provides statistically meaningful results in support of this claim. Moreover, we also cannot reject the hypothesis that the change in the HIBOR is equal to the change in the fed funds effective rate in response to monetary policy surprises. In addition, the adjusted R^2 for the HIBOR and the federal funds rate regressions are comparable once we focus on *changes* rather than *levels*. Finally, columns 4 and 6 suggest that the divergence between the HIBOR and the federal funds rate that started in late 2003 is not important for the effect of U.S. monetary policy surprises on Hong Kong's overnight rates.[18]

4 Stock Prices and Monetary Policy Shocks

4.1 The Severity of Omitted Variable Bias

In this section, we merge the econometric models presented in Section 2. Accordingly, we suppose that the target rate change, the U.S. stock price change, and the Hong Kong stock price change are

[17]The quarterly bulletin of HKMA attributes the spread in the Asian crisis period and the period between 2003-2005 to currency speculation, and the period thereafter to increased interbank liquidity and IPO waves, see http://www.hkma.gov.hk/media/eng/publication-and-research/quarterly-bulletin/qb200803/fa3_print.pdf. We check the robustness of our results by taking these periods into account.

[18]Here and henceforth, we determine outliers using the same criterion as in Bernanke and Kuttner (2005). Note that due to the time zone difference and holiday schedules, we do not have data from Hong Kong for each event date.

given by the following system:

$$\begin{aligned}
\Delta i &= \beta \Delta s + \gamma z + \varepsilon \\
\Delta s &= \alpha \Delta i + z + \eta \\
\Delta y &= a \Delta i + ez + w,
\end{aligned}$$

where z, ε, η, and w are orthogonal to each other. The first two equations capture the simultaneity and omitted variable problems through $\beta \neq 0$ and $\gamma \neq 0$. The third equation captures the possibility that Hong Kong stock returns can be affected by z, a vector of variables that also affect U.S. monetary policy and U.S. stocks. We can think of z and its coefficients as vectors, with γz and ez being the scalar products of parameter and variable vectors.

If we run the OLS regression of Δy on Δi, ignoring omitted variables, we get

$$\operatorname{plim} \hat{a}_{OLS} = \frac{cov\left(\Delta y, \Delta i\right)}{var\left(\Delta i\right)} = a + \frac{cov\left(ez, \Delta i\right)}{var\left(\Delta i\right)}.$$

So, unless $e = 0$, we have an omitted variable bias, although the regression does not suffer from a simultaneity problem.

How strong is this bias? When we run an instrumental variable regression of Δy on Δi, where the instrument is Δs, we get

$$\operatorname{plim} \hat{a}_{IV} = \frac{cov\left(\Delta s, \Delta y\right)}{cov\left(\Delta s, \Delta i\right)} = a + \frac{cov\left(\Delta s, ez\right)}{cov\left(\Delta s, \Delta i\right)},$$

which is equal to a if $e = 0$. This analysis implies that under the null hypothesis $e = 0$, meaning if we do not have an omitted variable bias, we should have $\operatorname{plim} \hat{a}_{IV} = \operatorname{plim} \hat{a}_{OLS} = a$ and $\hat{a}_{OLS}$ should be efficient. Therefore, we can test this hypothesis using the Hausman (1978) specification test.[19]

Table 2 reports the results from OLS regressions of the daily growth rate of the HSI on the expected and surprise funds target rate changes, and the same regression with the surprise rate change instrumented by the CRSP value-weighted equity return. Note that the coefficients on the surprise change under OLS and instrumental variable (IV) specifications are both statistically and quantitatively significantly different from each other. The difference persists even after the outliers are excluded. According to the above argument, the substantial difference between the coefficients serves as evidence that a potentially severe omitted variable bias exists if we specify our model as $\Delta y = a\Delta i + w$.[20]

4.2 Using U.S. Stock Returns to Control for Omitted Variable Bias

In this section we estimate the model

$$\Delta y = a\Delta i + b\Delta s + w$$

and provide evidence that this specification does not suffer from an omitted variable bias. We let t denote the event date. Note that the day before the event date, $t - 1$, does not include any target rate change by the Federal Reserve that may affect Hong Kong stock prices, due to the FOMC blackout period that is in effect before scheduled FOMC meetings.

[19] The conventional use of the Hausman test assumes that the inefficient (IV) estimator is consistent not only under the null hypothesis, but also under the alternative hypothesis. While $\hat{a}_{IV}$ above is consistent under the null hypothesis, $e = 0$, it does not have to be consistent if the null hypothesis is violated. Nevertheless, this does not pose a problem for the suggested test because, as stated by Cameron and Trivedi (2005, p. 273), ”the Hausman test is a quite general procedure that does not explicitly state an alternative hypothesis” and the properties of the Hausman test statistics are derived under the null hypothesis we aim to test.

[20] Similar to the estimates of Lucca and Moench (2012) for the U.S. economy, the OLS estimate shows a positive intercept. However, this intercept decreases in size and significance after controlling for outliers, and practically disappears after we control for omitted variables in the next section.

Using this information, we estimate the following model:

$$\Delta y_k = a\Delta i_k + (b + dc)\Delta s_k + w_k,$$

where $k \in \{t-1, t\}$ and d is a dummy variable equal to one at the pre-event dates and zero at event dates. Moreover, we take Δi_{t-1} to be zero at pre-event dates to capture the absence of a target rate change at pre-event dates.[21]

Under the null hypothesis that there are no omitted variables, the estimate for the coefficient of Δs should be the same on both the event and pre-event dates, that is, $c = 0$. To see the intuition, note that under the proposed econometric model we have

$$\begin{aligned}
\Delta i &= \beta\Delta s + \gamma z + \varepsilon \\
\Delta s &= \alpha\Delta i + z + \eta \\
\Delta y &= a\Delta i + b\Delta s + ez + w.
\end{aligned}$$

If $e \neq 0$, the simple OLS estimation of Δy on Δi and Δs (omitting the variables z) would lead to biased estimates of a and b because $cov\,(\Delta i, ez) \neq 0$ and $cov\,(\Delta s, ez) \neq 0$. Moreover, the magnitude of the bias for coefficient b would be different at event and pre-event dates if $e \neq 0$ because there are no monetary policy shocks at pre-event dates. Hence, we can conclude that $e = 0$ if the estimates of b for event and pre-event dates are the same, or equivalently if $c = 0$. Therefore, testing $c = 0$ in the proposed regression is the same as testing $e = 0$, meaning, for the omitted variable bias. The appendix presents this argument at a more formal level and also shows that the estimates of a and b obtained from this regression are unbiased when $e = 0$.[22]

[21] We choose the pre-event dates as our "control group" because the Fed has a strict blackout period before FOMC meetings.

[22] One can argue that there might be changes in investors' expectations regarding the outcome of FOMC delibera-

Based on the different responses at event and pre-event dates, Rigobon and Sack (2004) suggest using an instrumental variable model to control for endogeneity and omitted variable problems. Our approach differs from theirs because we offer a standard OLS model that directly accounts for omitted variables, and we then use event and pre-event observations to support the validity of our model. Moreover, unlike Rigobon and Sack (2004), our identification method does not assume that the nonmonetary shocks and variables are homoscedastic at event and pre-event dates.

The last line of Table 3 shows that we cannot reject the hypothesis $c = 0$ with or without the exclusion of outliers. This is consistent with our hypothesis that including U.S. stock returns as a regressor controls for omitted variable bias.

4.3 Robustness

An interesting result in Table 3 is that none of the dates in the Asian financial crisis are discarded as outliers, although these included a turbulent period for the Hong Kong economy that involved a speculative attack on the Hong Kong dollar. Table 6 in the appendix further shows that our results do not change significantly when we extend the time period to cover the entire 1989–2008 period and add control dummy variables for the major currency speculation period during the Asian crisis and for the period after September 2003 when we start to observe a gap between the HIBOR and the federal funds target rate.[23]

As a final robustness check, we also have added the change in the HSI index on the previous day as an additional variable in the regressions. After controlling for outliers, the estimated coefficient on this additional variable is economically and statistically insignificant, while the other coefficients have stayed essentially the same. Hence the results of this last regression are not reported here.

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tions at pre-event dates despite the black-out period. We repeat our analysis using the fed funds future price changes at pre-event dates for Δi_{t-1}, instead of taking $\Delta i_{t-1} = 0$. The appendix shows that the intuition presented here is still valid and the corresponding empirical results remain very similar.

[23]Major attacks occurred on October 1997 and January, June, and August 1998, so we interact the surprise term with a dummy that is equal to one between October 1997 and August 1998.

5 Malaysia: The Other Side of the Impossible Trinity

In this section, we further illustrate the power of our identification through using the Impossible Trinity by providing an example at the other side of the trinity. On September 30, 1998, the Malaysian government responded to the Asian financial crisis in an unorthodox way, compared to the responses by other East Asian countries. As shown in Figure 4, the Malaysian ringgit was pegged at 3.80 ringgits to the U.S. dollar, but foreign capital repatriated before staying at least 12 months became subject to substantial levies, and several limitations were imposed on bank and foreign transactions. The peg lasted until July 21, 2005, after which Malaysia switched to a managed floating exchange rate against an undisclosed basket of currencies while the capital controls were still in place.[24]

According to the Impossible Trinity theory, the combination of a fixed exchange rate and capital controls implies that Malaysia's monetary policy is independent of U.S. monetary policy.[25] Figure 5 seems to confirm this result: Both the policy rate of Bank Negara Malaysia and the interbank overnight rates in Malaysia (KLIBOR) are unaffected by changes in the federal funds target rate.[26] Therefore, if our identification strategy has merit, we should expect that Malaysian stock prices would barely respond to monetary policy shocks from the United States. Table 4 compares the response of the HSI and the Kuala Lumpur Composite Index (KLCI) to changes in the U.S. federal funds target rate. This table clearly fulfills our expectations, and hence supports our identification

[24] Several of these capital controls have been gradually relaxed. However, quite a few restrictions, such as limitations on foreign exchange transactions and payment of profits, dividends, and rental income to nonresidents, have survived at least until the end of the sample period examined in this section. A detailed history of Malaysian capital controls can be found in Johnson, Kochhar, Mitton, and Tamirisa (2007).

[25] Malaysia is a better example than other countries with the same properties, such as China, because it is a market economy. Moreover, Malaysia and Hong Kong have similar trade patterns with the United States, and their stock market indices have similar composition. For example, according to IMF DOTS/WEO 2004 data, the ratio of the total trade with the United States to local GDP is 35 percent for Hong Kong and 32 percent for Malaysia. Also, the largest 30 firms that constitute more than 80 percent of the value in Hong Kong and Malaysian stock indices include a similar number of banks and financial institutions: six in Malaysia and eight in Hong Kong as of 2004. This makes Malaysia a suitable alternative case study.

[26] During this time period, the Malaysian primary policy rate was the BNM 3-month intervention rate until the Overnight Policy Rate (OPR) was introduced as the policy rate in April 2004.

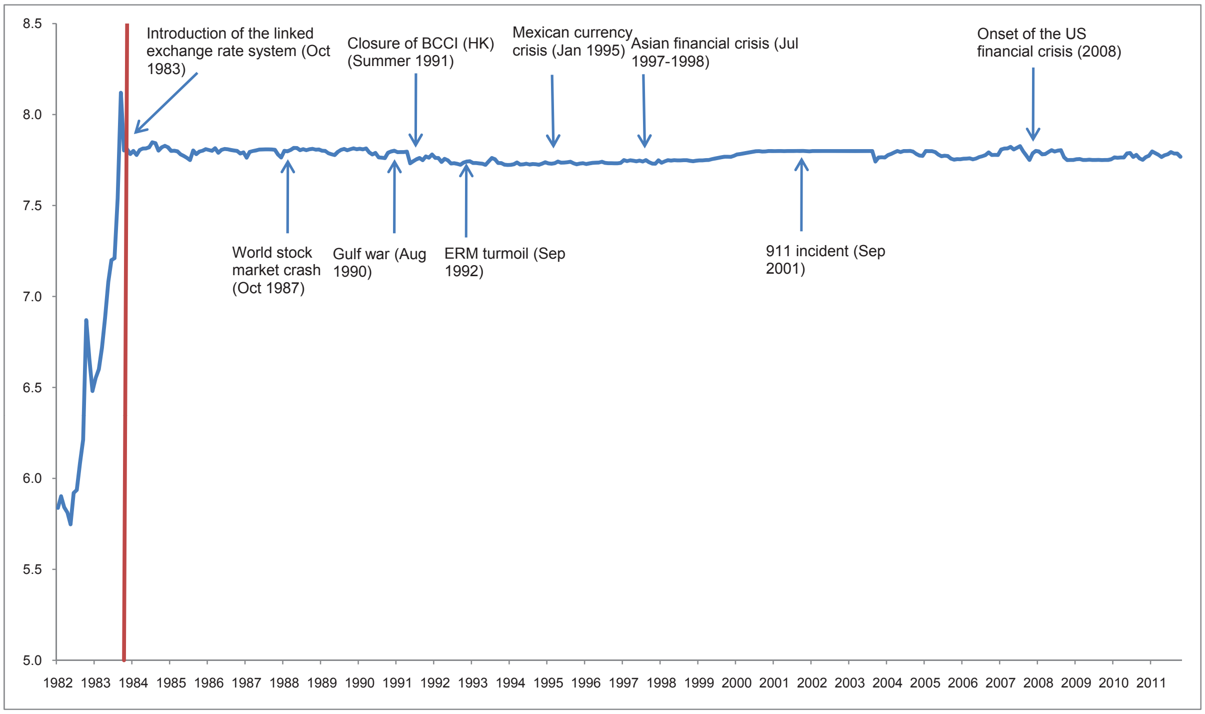
mechanism.

6 Conclusion

Basing our analysis on Mundell and Fleming's Impossible Trinity theory, we identify the impact of monetary policy on asset prices using Hong Kong stock market data and surprise changes in the U.S. federal funds target rate. As summarized in Rigobon and Sack (2004), two major problems arise in estimating the stock market's response to monetary policy. One is that monetary policy is simultaneously influenced by fluctuations in the stock market. The other is that there may be factors that have a direct impact both on monetary policy and the stock market, which creates an omitted variable bias. By focusing on the Hong Kong stock market's response to U.S. monetary policy, we circumvent the simultaneity problem, since changes in Hong Kong stock prices do not directly influence U.S. monetary policy. We also show that using U.S. stock returns as an additional regressor controls for omitted variable bias. Our results indicate that a 1 percent surprise increase in monetary policy rate leads to a 4 to 5 percent change in stock prices, which is significantly less than the 7 to 8 percent change implied by a traditional event study approach.

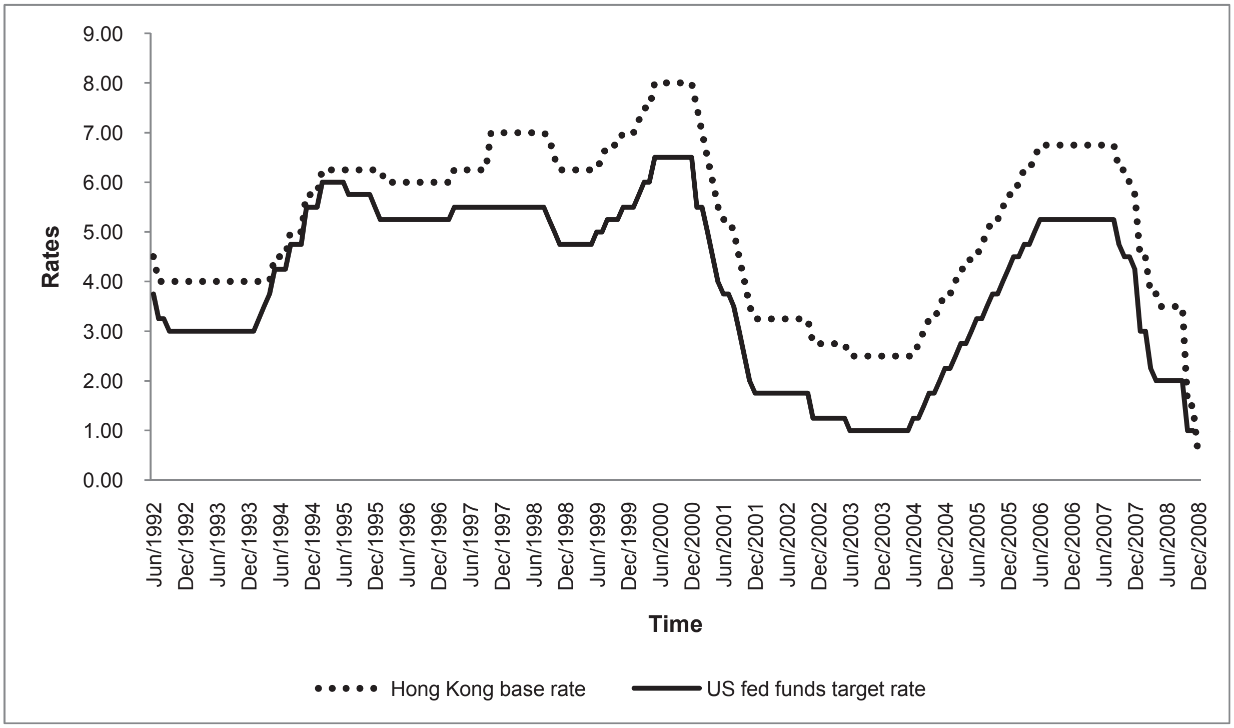
7 Figures and Tables

Figure 1: HK dollar/U.S. dollar Exchange Rate



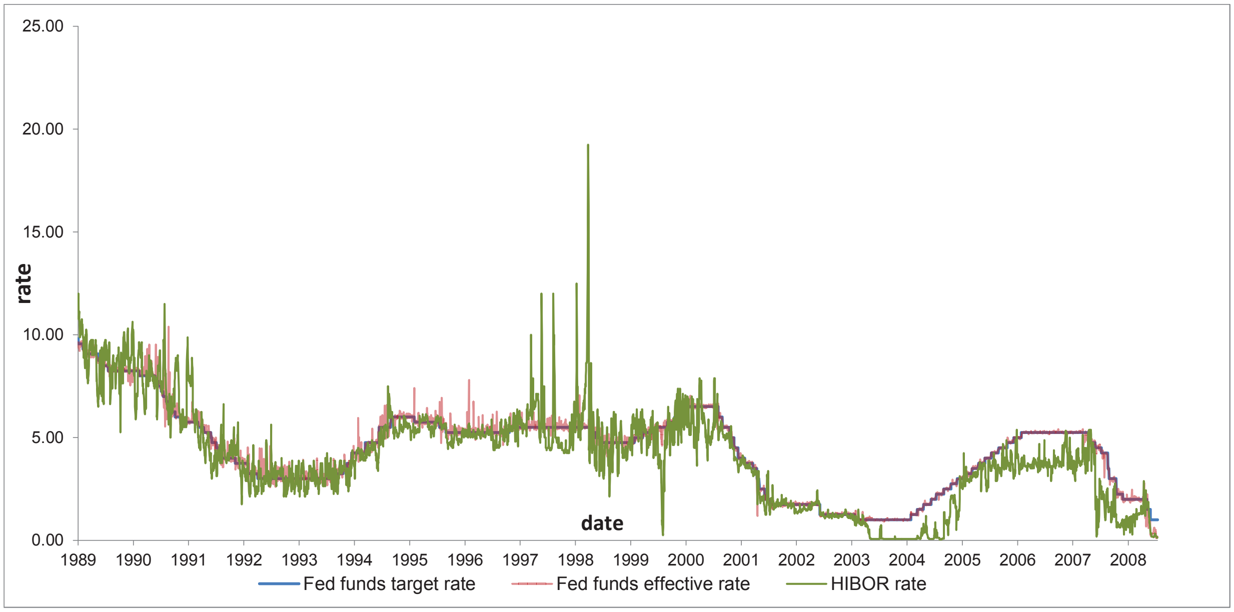
Source: This figure replicates the figure on p.36 of HKMA background brief No.1 (available at http://www.hkma.gov.hk/media/eng/publication-and-research/background-briefs/hkmalin/full_e.pdf), but extends the time period to 2011. Monthly data on the HKD/USD exchange rate can be downloaded from Bloomberg (ticker: HKD CURNCY).

Figure 2: Hong Kong Base Rate vs. U.S. Federal Funds Target Rate (End of Month)



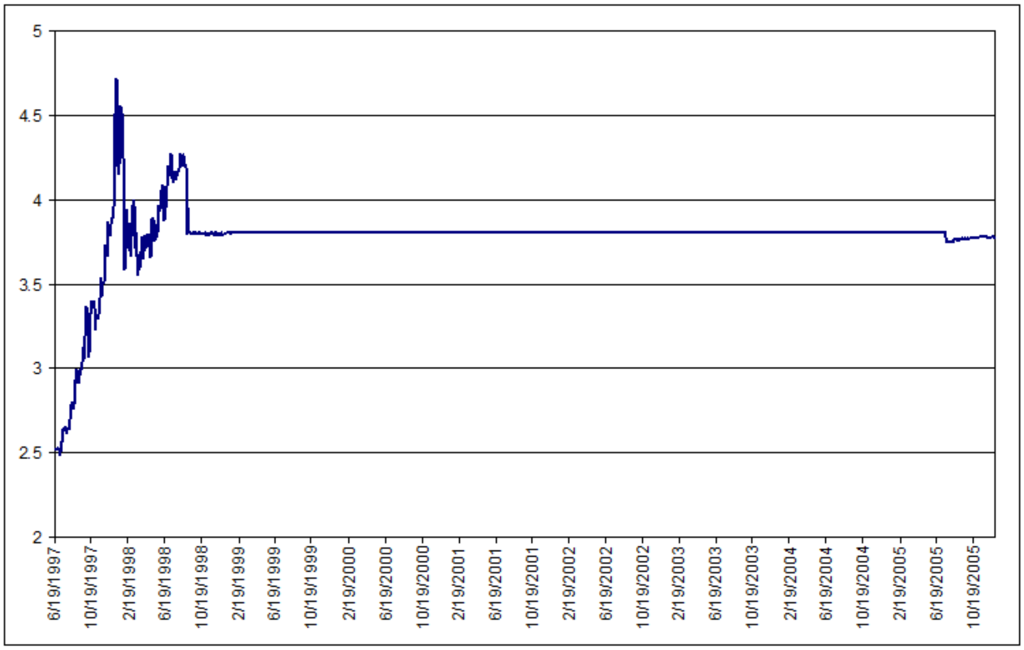
Source: End-of-month data on the Hong Kong base rate and the U.S. federal funds target rate are available at http://www.info.gov.hk/hkma/eng/statistics/index_efdhk.htm and http://research.stlouisfed.org/fred2/series/DFEDTAR, respectively. The U.S. federal funds target rate data ends in 2008, when the Federal Reserve stopped announcing a specific target rate and started announcing a range.

Figure 3: HIBOR Rate and Fed Funds Effective Rate vs. Fed Funds Target Rate, 1989–2008



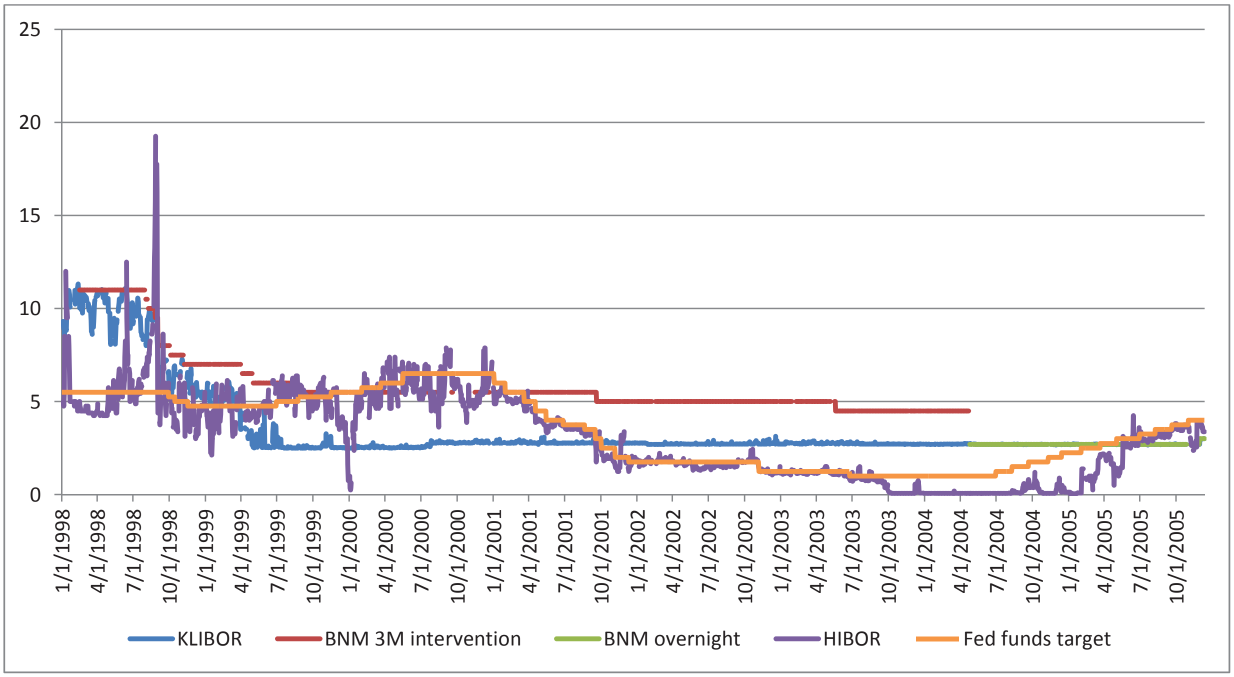
Source: The daily HIBOR rate is available at http://www.hkma.gov.hk/eng/market-data-and-statistics/monthly-statistical-bulletin/table.shtml#section5 (Section 6.3.3). The daily effective federal funds rate and fed funds target rate are available at http://research.stlouisfed.org/fred2/series/DFF and http://research.stlouisfed.org/fred2/series/DFEDTAR, respectively. The U.S. federal funds target rate data ended in 2008 when the Federal Reserve ceased announcing a specific target rate and started announcing a range.

Figure 4: Malaysian ringgit/U.S. dollar Exchange Rate, 1997–2005



Source: Bloomberg

Figure 5: Overnight Interest Rates in Malaysia and Hong Kong, 1998–2005



Source: Bloomberg

Table 1: HK Overnight HIBOR/ U.S. Federal Funds Effective Rate (Level/Change) vs. U.S. Federal Funds Target Rate (Level/Change), 1989–2008

	Level		Change					
	Full sample		Full sample			Excluding outliers		
Regressor	HIBOR	FF effective	HIBOR	HIBOR^p	FF effective	HIBOR	HIBOR^p	FF effective
Intercept	-0.66***	0.03	-0.07	-0.07	-0.02	-0.03	-0.03	-0.01
	(0.10)	(0.03)	(0.04)	(0.04)	(0.02)	(0.03)	(0.03)	(0.02)
FF target rate	1.05***	1.00***	-	-	-	-	-	-
	(0.02)	(0.01)	-	-	-	-	-	-
Expected change	-	-	0.28	0.29	0.24*	0.46***	0.46***	0.27**
	-	-	(0.20)	(0.20)	(0.13)	(0.13)	(0.13)	(0.12)
Surprise change	-	-	0.85***	0.76*	0.75***	1.18***	1.24***	0.74***
	-	-	(0.29)	(0.41)	(0.18)	(0.15)	(0.23)	(0.19)
Surprise $\times$ post-Sep03	-	-	-	0.26	-	-	-0.14	-
	-	-	-	(0.41)	-	-	(0.26)	-
Adjusted R^2	0.85	0.98	0.03	0.02	0.10	0.14	0.14	0.09
Obs.	323	323	152	152	152	149	149	149
χ^2: HIBOR = FF effective	3.67*		0.06	0.00	-	1.65	1.48	-

Note: Post-Sep03 is a dummy variable that takes the value one after September 30, 2003, and zero otherwise. in order to capture the period of divergence between HIBOR and the federal funds rate. Observations whose Cook's distance statistic exceeds 0.1 are considered as outliers.

$$Cook's\ d = \frac{\Delta\hat{\theta}_t^{'}\hat{\Sigma}^{-1}\Delta\hat{\theta}_t}{k},$$

where $\Delta\hat{\theta}_t$ is the change in the vector of regression coefficients resulting from dropping oberservation t, $\hat{\Sigma}$ is the estimated covariance matrix of the coefficients, and k is the number of regressors (including the constant) of the regression. There are no outliers for the HIBOR level and FF effective level regressions. The outliers for the HIBOR change regression are July 2, 1992, August 19, 1997, and May 16, 2000. For the sake of comparability, the outliers for the HIBOR^p and FF effective change regressions are the same as those for the HIBOR change regression, where the superscript p denotes that post-2003 dummy is included. The last row of this table reports χ^2 obtained from the post-estimation of the seemingly unrelated regression (SUR) system consisting of HIBOR (level/change) and FF effective (level/change) equations. The first post-estimation test is on the coefficient "FF target rate". The other post-estimation tests are on the coefficient "Surprise change". Robust standard errors are reported in parentheses. The symbols ***, **, and * indicate significance levels at 1 percent, 5 percent, and 10 percent, respectively.

Table 2: The Response of the HK Equity Return to Federal Funds Rate Changes, 1994–2005

	Full sample		Excluding outliers	
Regressor	OLS	IV	OLS	IV
Intercept	0.31*	-0.11	0.22	0.10
	(0.17)	(0.46)	(0.16)	(0.16)
Expected change	1.12	2.84	0.39	0.81
	(0.89)	(2.39)	(0.75)	(0.93)
Surprise change	-7.94***	-29.07	-7.67***	-15.66***
	(2.78)	(18.76)	(1.54)	(3.98)
Adjusted R^2	0.15	-	0.16	-
Obs.	87	87	84	84
Hausman test (χ^2)	-	5.76*	-	5.69*
Robust Hausman (t)	-	3.64***	-	2.69***

Notes: In the IV regression, we use the U.S. equity return as an instrumental variable for surprise changes. Observations whose Cook's distance statistic exceeds 0.1 are considered as outliers.

$$\textit{Cook's } d = \frac{\Delta\hat{\theta}_t^{'}\hat{\Sigma}^{-1}\Delta\hat{\theta}_t}{k},$$

where $\Delta\hat{\theta}_t$ is the change in the vector of regression coefficients resulting from dropping oberservation t, $\hat{\Sigma}$ is the estimated covariance matrix of the coefficients, and k is the number of regressors (including the constant) of the regression. The outliers for the OLS regression are May 17, 1994, October 15, 1998, and September 17, 2001. For the sake of comparability, the outliers for the IV regression are the same as those for the OLS regression. Robust standard errors are reported in parentheses. The symbols ***, **, and * indicate significance levels at 1 percent, 5 percent, and 10 percent, respectively. The Hausman test is from Hausman (1978) where OLS is assumed to be efficient. The robust Hausman test is the two-stage test discussed in section 8.4.3 of Cameron and Trivedi (2005) which uses robust standard errors.

Table 3: The Response of HK Equity Returns to U.S. Federal Funds Rate Changes and U.S. Equity Returns, 1994–2005

	Full sample		Excluding outliers	
Regressor	1(a)	2(a)	1(b)	2(b)
Intercept	0.20	0.17	0.12	0.14
	(0.16)	(0.11)	(0.14)	(0.10)
Surprise change	-5.38***	-5.43***	-4.69***	-4.69***
	(1.88)	(1.85)	(1.74)	(1.72)
Expected change	0.43	0.44	0.00	0.00
	(0.90)	(0.90)	(0.79)	(0.78)
US equity returns	0.57***	0.57***	0.44***	0.44***
	(0.16)	(0.16)	(0.16)	(0.16)
US equity returns (pre-event)	-	-0.09	-	0.03
	-	(0.20)	-	(0.21)
Adjusted R^2	0.29	0.24	0.23	0.18
Obs.	87	168	84	163

Notes: Regressions 1(a) and 1(b) use observations only on event dates, and regressions 2(a) and 2(b) use observations on both event and pre-event dates. The number of observations for regression 1(b) is less than 174 (87×2) because there are missing variables in the event and pre-event dates sample. Observations whose Cook's distance statistics exceeds 0.1 are considered as outliers.

$$\textit{Cook's } d = \frac{\Delta\hat{\theta}_t^{'}\hat{\Sigma}^{-1}\Delta\hat{\theta}_t}{k},$$

where $\Delta\hat{\theta}_t$ is the change in the vector of regression coefficients resulting from dropping oberservation t, $\hat{\Sigma}$ is the estimated covariance matrix of the coefficients, and k is the number of regressors (including the constant) of the regression. The outliers for regression 1(b) are May 17, 1994, October 15, 1998, and September 17, 2001. The outliers for regression 2(b) are the outliers for regression 1(b) and their corresponding pre-event dates, namely May 16, 1994, May 17, 1994, October 14, 1998, October 15, 1998, and September 17, 2001. September 16, 2001 is not an outlier because it is not included in the event and pre-event dates regression due to missing variables. Robust standard errors are reported in parentheses. The symbols ***, **, and * indicate significance levels at 1 percent, 5 percent, and 10 percent, respectively.

Table 4: The Response of HSI and KLCI Stock Prices to U.S. Federal Funds Rate Changes.

	Full sample		Excluding outliers	
Regressor	HSI	KLCI	HSI	KLCI
Intercept	-0.21	0.20	-0.18	0.13
	(0.15)	(0.14)	(0.14)	(0.14)
Expected change	-0.54	-0.10	-0.27	0.75
	(0.94)	(0.83)	(0.79)	(0.66)
Surprise change	-7.78***	0.05	-5.88***	-0.15
	(1.89)	(0.74)	(1.04)	(0.67)
US equity returns	0.50***	0.06	0.41***	0.06
	(0.13)	(0.06)	(0.07)	(0.07)
Adjusted R^2	0.51	-0.05	0.38	-0.03
Obs.	51	51	46	46

Note: The data span from September 30, 1998, to July 21, 2005, period when Malaysia had a fixed exchange rate. The last day of the Fed's monetary policy action during this period was June 30, 2005. We choose the dates for which both KLCI and HSI data are available to maintain comparability. Observations whose Cook's distance statistics exceed 0.1 are considered as outliers.

$$\textit{Cook's } d = \frac{\Delta\hat{\theta}_t^{'}\hat{\Sigma}^{-1}\Delta\hat{\theta}_t}{k},$$

where $\Delta\hat{\theta}_t$ is the change in the vector of regression coefficients resulting from dropping oberservation t, $\hat{\Sigma}$ is the estimated covariance matrix of the coefficients, and k is the number of regressors (including the constant) of the regression. For the sake of comparability, the outliers are taken as the union of the set of outliers for the HSI and KLCI regressions. The outliers are October 15, 1998, May 16, 2000, January 3, 2001, March 20, 2001, and November 6, 2001. Robust standard errors are reported in parentheses. The symbols ***, **, and * indicate significance level at 1 percent, 5 percent, and 10 percent, respectively.

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9 Appendix: Using Pre-event Dates for Omitted Variable Test

Let t be the event date and $t-1$ be the pre-event date. In this section we show that if the true model is given by

$$\begin{aligned} \Delta y_{t-1} &= b\Delta s_{t-1} + ez_{t-1} + w_{t-1} \\ \Delta y_t &= a\Delta i_t + b\Delta s_t + ez_t + w_t, \end{aligned}$$

then testing the hypothesis $e = 0$ is equivalent to testing the hypothesis $c = 0$ in the following regression,

$$\Delta y = a(1-d)\Delta i + (b + cd)\,\Delta s + w,$$

where $d = 1$ for pre-event dates and zero otherwise. We do so by showing that $E\left(\hat{c}_{OLS}\right) = 0$ for the $\hat{c}_{OLS}$ that comes from this regression when $e = 0$.

Note that we can write this regression as

$$\begin{pmatrix} \Delta y_{t-1} \\ \Delta y_t \end{pmatrix} = \begin{pmatrix} 0 & \Delta s_{t-1} & \Delta s_{t-1} \\ \Delta i_t & \Delta s_t & 0 \end{pmatrix} \begin{pmatrix} a \\ b \\ c \end{pmatrix} + \begin{pmatrix} w_{t-1} \\ w_t \end{pmatrix},$$

where each variable gives a vector of observations. Then, the OLS estimates for the parameters

a, b, c are given by

$$\begin{pmatrix} \hat{a}_{OLS} \\ \hat{b}_{OLS} \\ \hat{c}_{OLS} \end{pmatrix} = \left[\begin{pmatrix} 0 & \Delta i_t' \\ \Delta s_{t-1}' & \Delta s_t' \\ \Delta s_{t-1}' & 0 \end{pmatrix} \begin{pmatrix} 0 & \Delta s_{t-1} & \Delta s_{t-1} \\ \Delta i_t & \Delta s_t & 0 \end{pmatrix} \right]^{-1} \begin{pmatrix} 0 & \Delta i_t' \\ \Delta s_{t-1}' & \Delta s_t' \\ \Delta s_{t-1}' & 0 \end{pmatrix}$$
$$* \begin{pmatrix} b\Delta s_{t-1} + ez_{t-1} + w_{t-1} \\ a\Delta i_t + b\Delta s_t + ez_t + w_t \end{pmatrix},$$

which leads to

$$\operatorname{plim} \hat{c}_{OLS} = \frac{cov\left(\Delta s_{t-1}, ez_{t-1}\right)}{var\left(\Delta s_{t-1}\right)} + \frac{cov\left(\Delta s_t, ez_t\right) var\left(\Delta i_t\right) - cov\left(\Delta i_t, ez_t\right) cov\left(\Delta i_t, \Delta s_t\right)}{cov\left(\Delta i_t, \Delta s_t\right)^2 - var\left(\Delta i_t\right) var\left(\Delta s_t\right)}.$$

Therefore, $\operatorname{plim} \hat{c}_{OLS} = 0$ iff $e = 0$. To be more precise, $\operatorname{plim} \hat{c}_{OLS} = 0$ is also satisfied by another condition that involves a non-linear restriction on model parameters. However, there is no economic justification for this restriction.[27] Therefore, we conclude that testing for $c = 0$ is equivalent to testing $e = 0$. Moreover, we do not need $var\left(z_t\right) = var\left(z_{t-1}\right)$ or $var\left(\eta_t\right) = var\left(\eta_{t-1}\right)$ for the validity of this test. Therefore, unlike Rigobon and Sack (2004), our identification mechanism does not need homoscedasticity of non-monetary shocks and variables.

Moreover, the OLS estimates are unbiased when $e = 0$. In particular, we have

$$\begin{aligned} \operatorname{plim} \hat{a}_{OLS} &= a + \frac{cov\left(\Delta s_t, ez_t\right) cov\left(\Delta i_t, \Delta s_t\right) - cov\left(\Delta i_t, ez_t\right) var\left(\Delta s_t\right)}{cov\left(\Delta i_t, \Delta s_t\right)^2 - var\left(\Delta i_t\right) var\left(\Delta s_t\right)} \\ \operatorname{plim} \hat{b}_{OLS} &= b - \frac{cov\left(\Delta s_t, ez_t\right) var\left(\Delta i_t\right) - cov\left(\Delta i_t, ez_t\right) cov\left(\Delta i_t, \Delta s_t\right)}{cov\left(\Delta i_t, \Delta s_t\right)^2 - var\left(\Delta i_t\right) var\left(\Delta s_t\right)} \end{aligned}$$

so that $\operatorname{plim} \hat{a}_{OLS} = a$ and $\operatorname{plim} \hat{b}_{OLS} = b$ if $e = 0$.

[27] The exclusion of nonlinear restrictions on model parameters is common in specification tests. For example, the Hausman test for endogeneity, also employed by Rigobon and Sack (2004) in the context of section 2.1, implicitly assumes that $1 - \alpha\beta \neq 0$.

9.1 What if $\Delta i_{t-1} \neq 0$?

In this case, we can write this regression as

$$\begin{pmatrix} \Delta y_{t-1} \\ \Delta y_t \end{pmatrix} = \begin{pmatrix} \Delta i_{t-1} & \Delta s_{t-1} & \Delta s_{t-1} \\ \Delta i_t & \Delta s_t & 0 \end{pmatrix} \begin{pmatrix} a \\ b \\ c \end{pmatrix} + \begin{pmatrix} w_{t-1} \\ w_t \end{pmatrix},$$

where each variable gives a vector of observation. Then, the OLS estimates for the parameters a, b, c are given by

$$\begin{pmatrix} \hat{a}_{OLS} \\ \hat{b}_{OLS} \\ \hat{c}_{OLS} \end{pmatrix} = \left[\begin{pmatrix} \Delta i'_{t-1} & \Delta i'_t \\ \Delta s'_{t-1} & \Delta s'_t \\ \Delta s'_{t-1} & 0 \end{pmatrix} \begin{pmatrix} \Delta i_{t-1} & \Delta s_{t-1} & \Delta s_{t-1} \\ \Delta i_t & \Delta s_t & 0 \end{pmatrix} \right]^{-1} \begin{pmatrix} \Delta i'_{t-1} & \Delta i'_t \\ \Delta s'_{t-1} & \Delta s'_t \\ \Delta s'_{t-1} & 0 \end{pmatrix}$$
$$* \begin{pmatrix} b\Delta s_{t-1} + ez_{t-1} + w_{t-1} \\ a\Delta i_t + b\Delta s_t + ez_t + w_t \end{pmatrix}.$$

This leads to

$$\begin{aligned} \operatorname{plim} \hat{c}_{OLS} = {} & \frac{\sigma_{i,s,t-1}\sigma_{i,s,t} + \sigma_{i,s,t}^2 - (\sigma_{i,t-1} + \sigma_{i,t})\,\sigma_{s,t}}{\sigma_{i,s,t}^2 \sigma_{s,t-1} + \left(\sigma_{i,s,t-1}^2 - (\sigma_{i,t-1} + \sigma_{i,t})\,\sigma_{s,t-1}\right)\sigma_{s,t}} \sigma_{s,ez,t-1} \\ & - \frac{\sigma_{i,s,t-1}\sigma_{i,s,t} + \sigma_{i,s,t-1}^2 - (\sigma_{i,t-1} + \sigma_{i,t})\,\sigma_{s,t-1}}{\sigma_{i,s,t}^2 \sigma_{s,t-1} + \left(\sigma_{i,s,t-1}^2 - (\sigma_{i,t-1} + \sigma_{i,t})\,\sigma_{s,t-1}\right)\sigma_{s,t}} \sigma_{s,ez,t} \\ & + \frac{\sigma_{i,s,t}\sigma_{s,t-1} - \sigma_{i,s,t-1}}{\sigma_{i,s,t}^2 \sigma_{s,t-1} + \left(\sigma_{i,s,t-1}^2 - (\sigma_{i,t-1} + \sigma_{i,t})\,\sigma_{s,t-1}\right)\sigma_{s,t}} \left(\sigma_{i,ez,t-1} + \sigma_{i,ez,t}\right), \end{aligned}$$

where $\sigma_{i,t} = var\,(\Delta i_t)\,, \sigma_{s,t} = var\,(\Delta s_t)\,, \sigma_{i,s,t} = cov\,(\Delta i_t, \Delta s_t)\,, \sigma_{i,ez,t} = cov\,(\Delta i_t, ez_t)$ and $\sigma_{s,ez,t} = cov\,(\Delta s_t, ez_t)$. Therefore, we have again the result that $p\lim \hat{c}_{OLS} = 0$ is satisfied if $e = 0$.

Moreover, the OLS estimates of a and b are unbiased when $e = 0$. In particular, we have

$$
\begin{aligned}
\operatorname{plim}\hat{a}_{OLS} &= a + \frac{\sigma_{i,s,t}\sigma_{i,s,t-1}\sigma_{s,ez,t} + \sigma_{i,s,t-1}\sigma_{s,t}\sigma_{s,ez,t-1} - \sigma_{s,t-1}\sigma_{s,t}\left(\sigma_{i,ez,t-1} + \sigma_{i,ez,t}\right)}{\sigma_{i,s,t}^2\sigma_{s,t-1} + \left(\sigma_{i,s,t-1}^2 - \left(\sigma_{i,t-1} + \sigma_{i,t}\right)\sigma_{s,t-1}\right)\sigma_{s,t}} \\
\operatorname{plim}\hat{b}_{OLS} &= b - \frac{\sigma_{i,s,t-1}\sigma_{i,s,t}\sigma_{s,ez,t-1} - \sigma_{i,s,t}\sigma_{s,t-1}\left(\sigma_{i,ez,t-1} + \sigma_{i,ez,t}\right)}{\sigma_{i,s,t}^2\sigma_{s,t-1} + \left(\sigma_{i,s,t-1}^2 - \left(\sigma_{i,t-1} + \sigma_{i,t}\right)\sigma_{s,t-1}\right)\sigma_{s,t}} \\
&\quad - \frac{\left(\sigma_{i,s,t-1}^2 - \left(\sigma_{i,t-1} + \sigma_{i,t}\right)\sigma_{s,t-1}\right)\sigma_{s,ez,t}}{\sigma_{i,s,t}^2\sigma_{s,t-1} + \left(\sigma_{i,s,t-1}^2 - \left(\sigma_{i,t-1} + \sigma_{i,t}\right)\sigma_{s,t-1}\right)\sigma_{s,t}}.
\end{aligned}
$$

In order to implement this regression, we calculate the Kuttner surprise for pre-event dates using the federal funds futures data to replace Δi_{t-1}. Because there is no announcement of a change in the federal funds target rate at pre-event dates, we do not have any measure for the expected component of the federal funds target change on these dates. Therefore, we also omit the expected component of fed funds target rate changes from event dates. This should not affect the results anyway since the expected component always turned out to be insignificant in our regressions.

The results, shown in Table 5, illustrate that $\hat{c}_{OLS}$ is both economically and statistically insignificant, providing evidence that $e = 0$. Also, note that if $e \neq 0$, the size of bias for these estimators, $\hat{a}$ and $\hat{b}$, should be significantly different from the size of the bias in our original estimator where we did not use the pre-event dates. We do not observe such a difference between the two estimators, a result which is also consistent with our hypothesis that $e = 0$.

Table 5: The Response of HK Equity Returns to U.S. Federal Funds Rate Changes and U.S. Equity Returns, 1994–2005.

	Full sample		Excluding outliers	
Regressor	1(a)	2(a)	1(b)	2(b)
Intercept	0.20	0.19*	0.12	0.16
	(0.16)	(0.11)	(0.14)	(0.10)
Unexpected change	-5.17**	-4.39**	-4.69***	-3.61**
	(1.99)	(1.88)	(1.71)	(1.63)
US equity returns	0.58***	0.60***	0.44***	0.47***
	(0.16)	(0.17)	(0.16)	(0.16)
US equity returns (pre-event)	-	-0.09	-	0.02
	-	(0.21)	-	(0.21)
Adjusted R^2	0.29	0.24	0.24	0.19
Obs.	87	168	84	163

Note: The policy shocks on pre-event dates are calcuated in the same way as Kuttner's surprises. Regressions 1(a) and 1(b) use observations only on event dates, and regressions 2(a) and 2(b) use observations on both event and pre-event dates. The number of observations for regression 1(b) is less than 174 (87×2) because there are missing variables in the event and pre-event dates sample. Observations whose cook's distance statistics exceeds 0.1 are considered as outliers.

$$\textit{Cook's } d = \frac{\Delta\hat{\theta}_t^{'}\hat{\Sigma}^{-1}\Delta\hat{\theta}_t}{k},$$

where $\Delta\hat{\theta}_t$ is the change in the vector of regression coefficients resulting from dropping oberservation t, $\hat{\Sigma}$ is the estimated covariance matrix of the coefficients, and k is the number of regressors (including the constant) of the regression. The outliers for regression 1(b) are May 17, 1994, October 15, 1998, and September 17, 2001. The outliers for regression 2(b) are the outliers for regression 1(b) and their corresponding pre-event dates, namely May 16, 1994, May 17, 1994, October 14, 1998, October 15, 1998, and September 17, 2001. September 16, 2001 is not an outlier because it is not included in the event and pre-event dates regression due to missing variables. Robust standard errors are reported in parentheses. The symbols ***, **, and * indicate significance levels at 1 percent, 5 percent, and 10 percent, respectively. The data to calculate the policy surprises at pre-event dates comes from Bloomberg, with the exception of November 14, 1994 and December 19, 1994 due to roundoff errors in Bloomberg that is evident from the mismatch with Kuttner surprises on event dates. For these dates we use Wikiposit/Futures-Data which matches Kuttner surprises.

Table 6: Asian Crisis and HIBOR Gap, 1989–2008

	Full sample				Excluding outliers			
Regressor	1(a)	2(a)	3(a)	4(a)	1(b)	2(b)	3(b)	4(b)
Intercept	0.17	0.16	0.18	0.17	0.16	0.14	0.16	0.14
	(0.12)	(0.12)	(0.12)	(0.12)	(0.11)	(0.11)	(0.11)	(0.11)
Expected change	0.49	0.47	0.20	0.17	-0.03	-0.05	-0.05	-0.08
	(0.69)	(0.69)	(0.68)	(0.68)	(0.54)	(0.54)	(0.56)	(0.56)
Surprise change	-7.34***	-7.31***	-3.68**	-3.60**	-3.53***	-3.47***	-3.44***	-3.36**
	(2.71)	(2.74)	(1.66)	(1.67)	(1.20)	(1.21)	(1.29)	(1.30)
Surprise $\times$ Asian crisis	-	-25.55	-	-32.25	-	-30.11	-	-30.36
	-	(25.59)	-	(25.87)	-	(25.24)	-	(25.36)
Surprise $\times$ HIBOR gap	-	-	-10.34***	-10.48***	-	-	-1.57	-1.76
	-	-	(1.99)	(2.00)	-	-	(4.71)	(4.61)
US equity returns	0.61***	0.62***	0.71***	0.72***	0.59***	0.61***	0.60***	0.61***
	(0.14)	(0.14)	(0.14)	(0.14)	(0.10)	(0.10)	(0.11)	(0.11)
Adjusted R^2	0.35	0.35	0.43	0.43	0.26	0.27	0.26	0.26
Obs.	153	153	153	153	150	150	150	150

Note: The Asian crisis dummy is set to one for observations during the Asian crisis from October 1, 1997 to August 30, 1998. The HIBOR gap dummy is set to one for observations after September 30, 2003, when the HIBOR rate became more than 100 basis point below the federal funds rate. Observations whose cook's distance statistics exceeds 0.1 are considered as outliers.

$$Cook's\ d = \frac{\Delta\hat{\theta}_t^{'}\hat{\Sigma}^{-1}\Delta\hat{\theta}_t}{k},$$

where $\Delta\hat{\theta}_t$ is the change in the vector of regression coefficients resulting from dropping oberservation t, $\hat{\Sigma}$ is the estimated covariance matrix of the coefficients, and k is the number of regressors (including the constant) of the regression. For the sake of comparability, the outliers for regressions 2(b), 3(b), and 4(b) are the same as those for regression 1(b), namely May 17, 1994, October 15, 1998, and January 22, 2008. Robust standard errors are reported in parentheses. The symbols ***, **, and * indicate significance levels at 1 percent, 5 percent, and 10 percent, respectively.