Monetary Policy Shocks and Stock Returns: Identification Through Impossible Trinity

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February 2012

Abstract

This paper aims to identify the effect of monetary policy shocks on stock prices through the lens of Mundell and Fleming's "Impossible Trinity" theory. Our identification strategy seeks to solve the simultaneity and omitted variable problems inherent in studies that focus on the effect of monetary policy on asset prices. Moreover, we use our identification strategy to test the hypothesis that stock prices of financially constrained firms are more responsive to monetary policy shocks. Our results so far do not support this hypothesis, which seems to contradict the financial accelerator theory presented in Bernanke, Gertler, and Gilchrist (1999) but is consistent with Lamont, Polk, and Saá-Requejo (2001) who find that the relative stock market performance of constrained firms does not reflect monetary policy or credit conditions.

JEL Codes: E4 (E44), E5 (E52, E58), G1 (G12, G15, G18), G3 (G32, G38)

Keywords: Stock prices, monetary policy, financial frictions, financial accelerator, 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 at the same time 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 which would cause a bias even if there is no 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 (absence of capital controls), and an independent monetary policy. Hong Kong is a clear example of Mundell and Fleming's theory. First, there are no restrictions on capital flows or on trading of financial assets in Hong Kong. Second, as shown in Figure 1, Hong Kong monetary authority (HKMA) has successfully implemented a fixed exchange rate for HKD/USD since October 1983.[3] Since the establishment of the Exchange Rate Link, the Hong Kong dollar exchange rate has remained stable in the face of various shocks. It remained unaffected by the 1987 stock market crash, the Gulf War in 1990, the

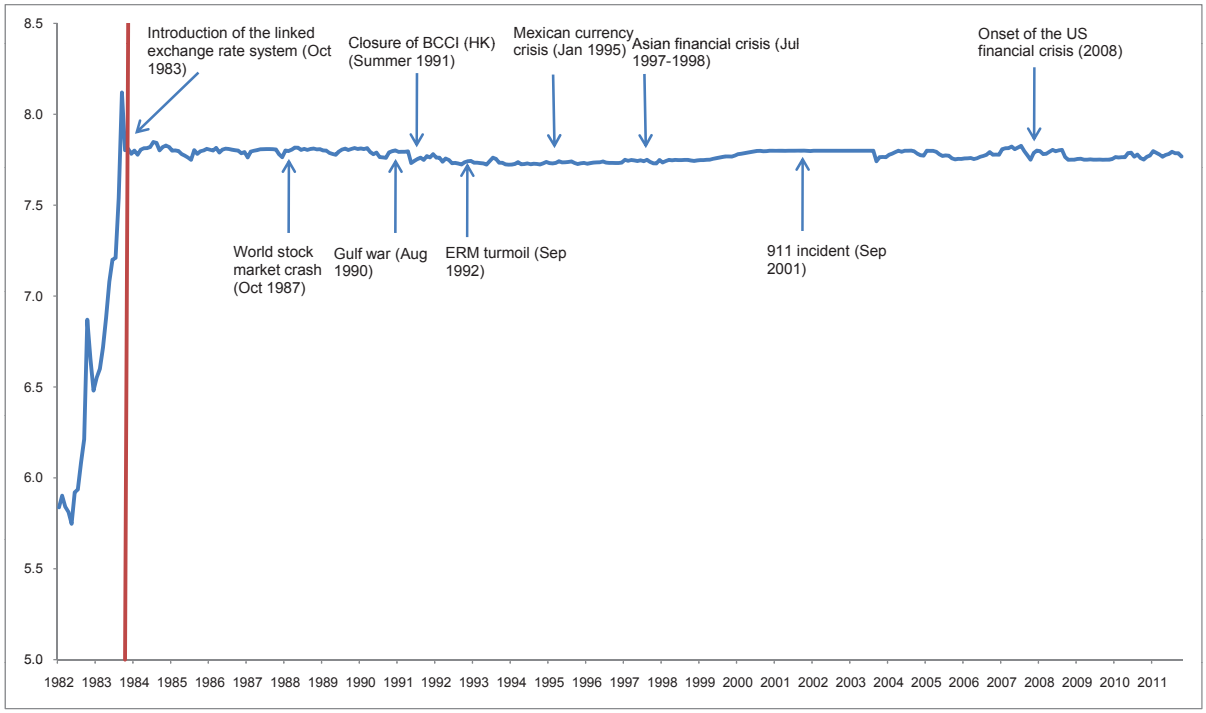
[1] See Rigobon and Sack (2003) for evidence of endogeneity problem.

[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 USD into HKD, or vice versa, at a predetermined rate until 2005. Since May 18, 2005 HKMA set a very narrow band of 7.85 as an upper limit and 7.75 as a lower limit for the HKD to flow within.

Exchange Rate Mechanism turmoil in Europe in 1992, the Mexican currency crisis of 1994/95, the Asian financial crisis of 1997/98, the 911 incident, and the 2008 crisis.

Figure 1: HKD/USD exchange rate



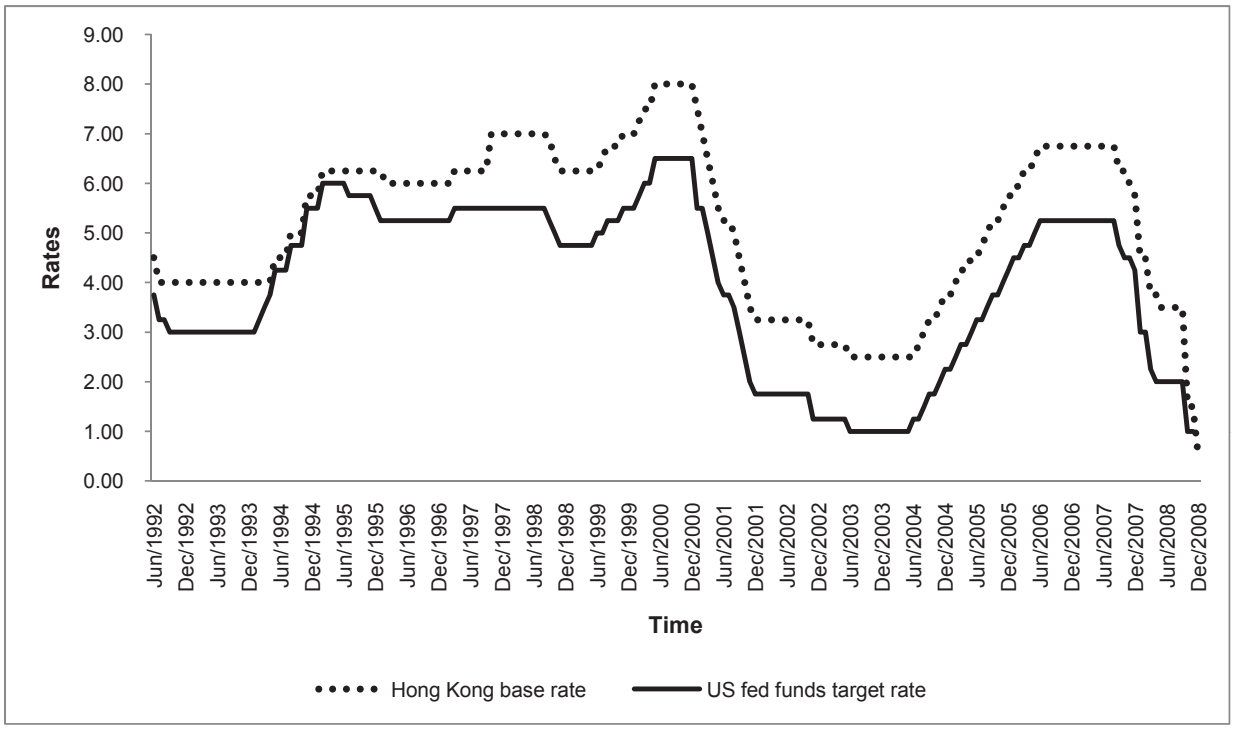
Note: This figure replicates the figure on p.36 of HKMA background brief No.1 (available at http://www.hkma.gov.hk/media/eng/publication-and-research/background-briefs/hkmalin/full_e.pdf), but extends the time period to 2011. Monthly data on HKD/USD exchange rate can be downloaded from Bloomberg (ticker: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

As a result, the Impossible Trinity suggests that the monetary policy of Hong Kong depends on US monetary policy. Figure 2 provides evidence for the close relationship between Hong Kong and US monetary policy - movements in Hong Kong base rate closely follows movements in fed funds target rate.[4] Given that changes in Hong Kong base rate closely follows changes in fed funds target rate and that U.S. Federal Reserve Bank does not base its monetary policy on the stock price

[4]Readers may be aware of the fact that the base rate is set mechanically by Hong Kong monetary authority through a transparent formula which is US 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*.

movements in Hong Kong explicitly, we can conclude that unexpected changes in fed funds target rate are exogenous shocks to Hong Kong monetary policy.[5]

Figure 2: Hong Kong base rate vs. US fed funds target rate (end of month)



Note: End-of-month data on Hong Kong base rate and US fed 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 US fed funds target rate data stops in 2008 when the Federal Reserve stops announcing a specific target rate and starts announcing a range.

To be sure, there have been episodes, in particular during the Asian financial crisis, where the markets have tested the ability of the Hong Kong Monetary Authority (HKMA) to stick to the fixed exchange rate. However, HKMA has managed to come out of these episodes victoriously. Moreover, the policies of the HKMA during the Asian crisis were mostly independent of the FOMC

[5]The Impossible Trinity can also be applied in the context of other countries that has 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 the peg is not a hard but an adjustable one 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 Bermuda dollar is at par with US dollar, but the stock market is not well-developed.

decisions except when HKMA adjusted the base rate to reflect the changes in the federal funds target rate.[6] Therefore, the federal funds target rate changes during this period can still be considered as exogenous monetary policy shocks to Hong Kong economy. Nevertheless, we control for these time periods in our robustness checks to confirm the stability of our results.

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

Finally, we use our identification strategy to study the relationship between financial frictions, stock prices and monetary policy shocks.[9] Using the financial accelerator framework of Bernanke, Gertler and Gilchrist (1999), we initially reveal a new implication of the credit channel of monetary

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

[7]By addressing the omitted variables issue explicitly, we allow the possibility that US monetary policy responds to macroeconomic events that affect both US economy and Hong Kong economy directly. Therefore, our identification mechanism only precludes 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 but they do not find a similar statistically significant bias for the response of the stock prices.

[9]Hubbard (1995) and Bernanke, Gertler, and Gilchrist (1996) list three empirical implications of the broad credit channel of monetary policy: (i) external finance is more expensive for borrowers than internal finance due to agency (monitoring) costs; (ii) the cost gap between internal and external finance depends inversely on the borrower's net worth; (iii) adverse shocks to net worth should reduce borrowers' access to finance, thereby reducing their investment, employment, and production levels. These implications are extensively studied in the literature, for example by Gertler and Gilchrist (1994), Kashyap, Lamont, and Stein (1994), and Oliner and Rudebusch (1996) among others.

policy: The stock prices of financially more constrained firms are more responsive to monetary policy shocks. Then, we use our framework to search for evidence of this hypothesis in the data.

Because the financial constraints in Bernanke, Gertler, and Gilchrist (1999) stem from monitoring costs, we first analyze how monetary policy affects the prices of stocks that are cross-listed in Hong Kong and U.S. in comparison to stocks that are listed only in Hong Kong, following the large body of literature that shows cross-listing of non-US firms in US reduces monitoring costs.[10] As a second test, we follow Gertler and Gilchrist (1994) who argue that ”in nearly every study the ”likely to be constrained” firms are much smaller on average than the control group” and use size as a proxy for financial constraint. Neither of these tests supports that more constrained firms' stock prices react more strongly to monetary policy shocks. While these results are consistent with the findings of Lamont, Polk, and Saá-Requejo (2001) they do seem to contradict the financial accelerator hypothesis, an issue which warrants further attention.

Our paper also 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 announcement, thereby distinguishing the impact of the policy change from the effects of news arriving earlier or later in the day. Bjørland and Leitemo (2009) use both short-run and long-run restrictions in a VAR framework to control for endogeneity.[11]

Our empirical analysis is closely related to Bernanke and Kuttner (2005), who study the reaction of the US stock market to federal funds target rate changes. Following their study, we use changes in federal funds futures' price on the dates of monetary policy announcements in order to

[10] This is called the ”bonding” theory. See, for example, Coffee (1999, 2002), Stulz (1999), Reese and Weisbach (2002), and Doidge, Karolyi, and Stulz (2004).

[11] These papers have a long line of predecessors that look at the relationship between monetary policy and asset prices and address the identification problems to various degrees, such as 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 (2005). See also Selin (2001) for an earlier survey.

identify surprise changes in 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.[12] However, we use Hong Kong stock returns on these event dates, rather than US 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 addresses this bias explicitly.

Our method also has other advantages in comparison to previous studies. Unlike Rigobon and Sack (2004), our identification method does not assume that non-monetary shocks and variables are homoscedastic.[13] Moreover, our identification mechanism allows us to omit some potential pitfalls of high frequency data. First, higher volatility of high frequency data can cause an amplified errors-in-variables. Second, the high frequency stock price drift prior to FOMC announcement, documented by Moench and Lucca (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, we show evidence that smaller stocks react to monetary policy shocks with a delay. This delayed response is particularly important for the cross-sectional study we are aiming in the second part of our paper. Fourth, even if the stock and fed funds futures market would react to FOMC related news immediately we can still have problems with identification if there have been other important macroeconomic news earlier in the day to which stocks don't react immediately, as explained below.

The studies that employ high frequency data assume that a narrow window, for example 15

[12] See Evans (1998) and Gürkaynak, Sack and Swanson (2007). Another advantage of looking at one-day changes in near-dated fed 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).

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

minutes before and 45 minutes after the announcement, is enough to capture the full effect of the FOMC announcement because the markets react to FOMC announcements very fast. Suppose that there has been another macroeconomic announcement earlier in the day, such as unemployment or CPI reports, to which markets don't react immediately. Even if the markets react to FOMC announcement immediately, a delayed response to other macroeconomic news implies that the movements in SP500 during the (-15m,+45m) event window is not only due to FOMC announcement but also partially due to delayed response to these other macroeconomic news. This will bring the identification problems back into picture: To the extent that the unemployment report has been incorporated in the FOMC decision but not yet in stock prices, the endogeneity problem is back in the picture because the response of stocks during the (-15m,+45m) event window is a superposition of the responses to FOMC announcement and unemployment report, although the FOMC announcement already incorporated the unemployment report and the associated delayed response of stocks. Also, to the extent that both the Fed funds futures and stock markets continue to respond to the macroeconomic news earlier in the day during the (-15m,+45m) event window, the omitted variables problem is back in the picture because part of the correlation between stock price and fed funds futures price in the event window is attributable to their joint response to unemployment report. It is not easily verifiable 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

The monetary policy might respond to stock returns at the same time as the stock returns respond to monetary policy. Suppose Δs is the change in US stock price and Δi is the change in the federal funds target rate. Then we have the standard simultaneous equation problem,

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

If we use 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 US monetary policy decisions directly and the changes in US monetary policy can be considered as exogenous shocks to Hong Kong economy according to the Mundell-Fleming model. That is, if we let Δy be stock price increase in Hong Kong we have

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

with $cov\left(\Delta i, w\right) = 0$. 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 decisions they might be indirectly correlated with these decisions when there are global shocks affecting both. This is what we focus on next.

2.2 Omitted Variable Problem

Some economic news that affect monetary policy might also affect stock prices directly in addition to their indirect effect through monetary policy. This 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 captures 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

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

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

Using Hong Kong stock market data does not address this problem directly. In particular, if there are any variables that affect both US 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 give 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 US 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 US stock returns as an additional regressor can eliminate it.

3 Data

The data for our empirical study fall into two categories: indices of US and Hong Kong equity market and variables that represent US monetary policy changes.

As in Bernanke and Kuttner (2005), we use total return on the CRSP value-weighted index as a measure of US equity return.[14] Our major indicator that keeps track of the stock market performance in Hong Kong is the daily Hang Seng index (HSI).[15]

One problem associated with the estimation of the market's reaction to monetary policy changes is that the 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 anticipated component of a monetary policy change, specifically, a change in federal funds target rate. Identification of the surprise element in target rate change relies on the price of 30-day federal funds futures contracts, 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 target rate by first calculating the change in the rate implied by the corresponding futures contract, given by 100 minus the future contract price, and then scaling it by a factor associated with the number of days of the month in which the event happens. Accordingly, the unanticipated target rate change, for an event taking place on day d of

[14]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 calculation methodology for index return.

[15]The Hang Seng index can be accessed through Bloomberg. Its ticker is HSI Index. We use the growth of HSI index in US dollars in our regressions to make the results comparable to studies that focus on US stock market. Using HSI index in HK dollars 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 secondary market exchange rate is not significantly correlated with US monetary policy surprises. Hong Kong stock market is closed at the time of scheduled FOMC announcements, since Hong Kong local time is twelve hours ahead of US eastern daylight saving time. We adjust forward one day for Hong Kong data.

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

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

where Δi is the actual funds rate change.

The data for the decomposition of fed funds target rate changes can be obtained from Kenneth Kuttner's webpage.[16] Kuttner's dataset contains futures-based funds rate surprise 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 target rate. In our initial analysis of stock prices, we focus primarily on the period between February 1994 to May 2005 for three reasons. First, starting February 1994, the policy of announcing target rate changes at pre-scheduled dates virtually eliminates the timing ambiguity associated with rate changes prior to this time period.[17] Second, Hong Kong has switched to a narrow floating band policy on May 18, 2005. Third, the same Federal Reserve governor, Alan Greenspan, has been in charge of monetary policy during this time period which decreases the contamination of our results by a potential change in policy regime.[18] We also check the robustness of our results by extending the dataset from June 1989 to June 2008.

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

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

[18] Alan Greenspan is succeeded by Ben Bernanke on February 1, 2006.

3.1 HIBOR versus Federal Funds rate

Before we start our analysis with the stock index, we want to provide further evidence regarding the close relationship between US monetary policy and the overnight interest rates in Hong Kong. Figure 3 presents overnight Hong Kong Interbank Offered Rate (HIBOR) which is the closest interest rate to federal funds effective rate. While HIBOR closely follows federal funds target rate, its track record is not as well as the federal funds effective rate, in particular during the Asian financial crisis of late 1990s and after September 2003. This observation is also confirmed by the comparison of adjusted R-squares in the left panel of Table 1. This pattern is expected because the banks in Hong Kong do not have direct access to the Federal Reserve facilities as banks in the US do.[19]

Nevertheless, for our identification mechanism to hold, i.e., for federal funds target rate changes to be considered as exogenous monetary policy shocks to Hong Kong economy, it is enough that a surprise change in federal funds target rate causes a proportionate change in HIBOR rate. The right panel of Table 1 provides results in support of this claim which we cannot reject statistically. Moreover, we also cannot reject that the change in HIBOR is equal to the change in fed funds effective rate in response to monetary policy surprises. In addition, adjusted R^2 for HIBOR and fed funds rate regressions are comparable once we focus on changes, rather than levels. Finally, columns 4 and 6 suggest that the divergence between HIBOR and fed funds rate that started in late 2003 is not important for the effect of US monetary policy surprises on Hong Kong overnight rates.[20]

[19]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.

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

Table 1: HK overnight HIBOR/ US federal funds effective rate (level/change) vs. US federal funds target rate (level/change) (1989-2008)

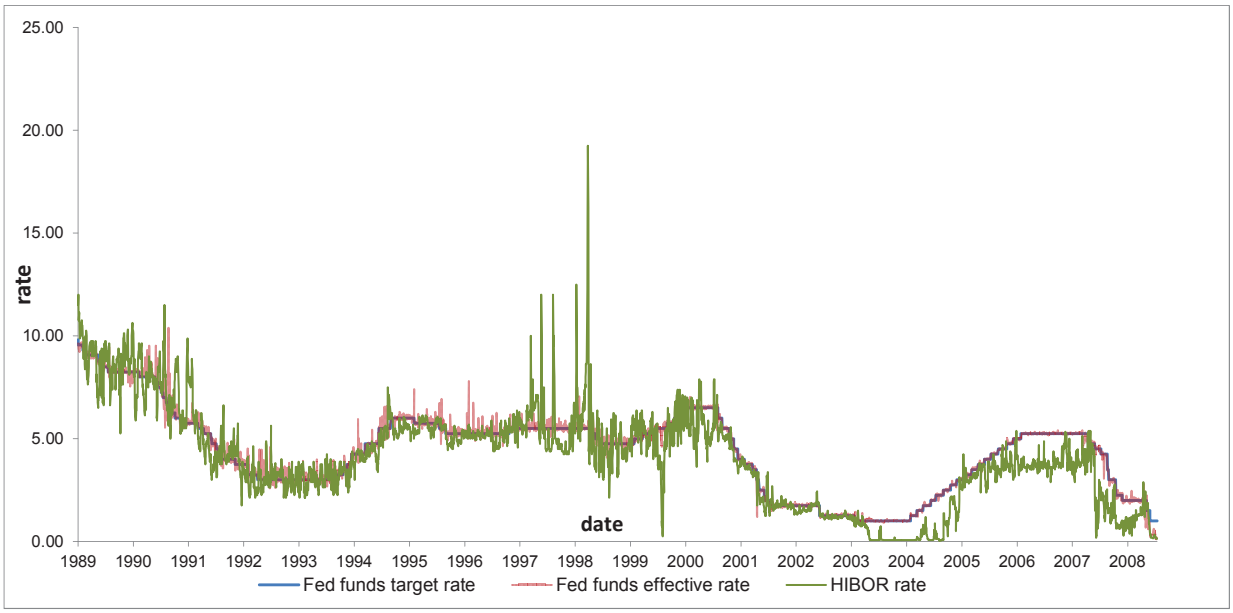
	Level		Change					
	Full sample		Full sample			Excluding outliers		
Regressor	HIBOR	FF effective	HIBOR	HIBOR^p	FF effective	HIBOR	HIBOR^p	FF effective
Intercept	-0.66***	0.03	-0.07	-0.07	-0.02	-0.03	-0.03	-0.01
	(0.10)	(0.03)	(0.04)	(0.04)	(0.02)	(0.03)	(0.03)	(0.02)
FF target rate	1.05***	1.00***	-	-	-	-	-	-
	(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 1 after September 30, 2003 and zero otherwise in order to capture the period of divergence between HIBOR and fed funds rate. 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. 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 HIBOR^p and FF effective change regressions are the same as those for HIBOR change regression. 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. ***, **, and * indicate significance level at 1%, 5%, and 10%, respectively.

Figure 3: HIBOR rate and fed funds effective rate vs. fed funds target rate



Note: 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). Daily effective fed 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 US fed funds target rate data stops in 2008 when the Federal Reserve stops announcing a specific target rate and starts announcing a range.

4 Stock Prices and Monetary Policy Shocks

4.1 Severity of omitted variable bias

In this section, we will merge the econometric models presented in Section 2. Accordingly, we suppose that the target rate change, US stock price change, and Hong Kong stock price change are 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, w are orthogonal to eachother. 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 some variables, z, that affect US monetary policy and US stocks. We can think of z and its coefficients as vectors, then γz and ez would be the scalar product of parameter and variable vectors.

If we run the OLS 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 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 $e = 0$, that is, if we do not have 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 Hausman (1978) specification test.[21]

Table 2 reports the results from OLS regressions of daily growth rate of Hang Seng index on the expected and surprise funds target rate changes, and the same regression with surprise rate change instrumented by CRSP value-weighted equity return. Note that the coefficients on surprise change under ordinary least squares (OLS) and instrumental variable (IV) specifications are both statistically and quantitatively significantly different from eachother. The difference persists even after the outliers are excluded. According to the argument above, this substantial difference serves as a piece of evidence that there exists a potentially severe omitted variable bias if we specify our model as $\Delta y = a\Delta i + w$.[22]

4.2 Using US stock returns to control for omitted variable bias

In this section we estimate the model

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

[21]The conventional use of 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.

[22]Similar to the estimates of Moench and Lucca (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.

Table 2: The response of 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***

Note: In the IV regression, we use US equity return as an instrumental variable for surprise changes. 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 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. ***, **, and * indicate significance level at 1%, 5%, and 10%, 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.

and provide evidence that this specification does not suffer from 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 Bank that may affect Hong Kong stock prices, due to FOMC blackout period.

Using this information, we estimate the 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.[23]

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, i.e., $c = 0$. 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 $e \neq 0$, 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 $e = 0$, i.e., for 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 $e = 0$.[24]

[23] We choose the pre-event dates as our 'control group' because the Fed has a strict black-out period before FOMC meetings.

[24] One can argue that there might be changes in investors' expectations regarding the outcome of FOMC at pre-event

Based on the different responses at event and pre-event dates, Rigobon and Sack (2004) suggest an instrumental variable model to control for endogeneity and omitted variable problems. Our approach differs from theirs because we offer a standard ordinary least squares model that directly takes omitted variables into account and we then use event and pre-event observations to argue for the validity of our model. Moreover, unlike Rigobon and Sack (2004), our identification method does not assume that the non-monetary 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, thereby concluding in favor of our hypothesis that including US 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 an outlier although this was a turbulent period for Hong Kong economy that involves a speculative attack to Hong Kong dollar. Table 8 in the appendix further shows that our results do not change significantly when we extend the time period in order to cover the whole 1989-2008 period and add control dummies for the major currency speculation period during the Asian crisis and for the period after September 2003 where we start to observe a gap between HIBOR and fed funds rate.[25]

As a final robustness check, we have also added the change in HSI index on previous day as an additional variable into the regressions. After controlling for outliers, the coefficient of 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.

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.

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

Table 3: The response of HK equity returns to US federal funds rate changes and US 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

Note: 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.

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

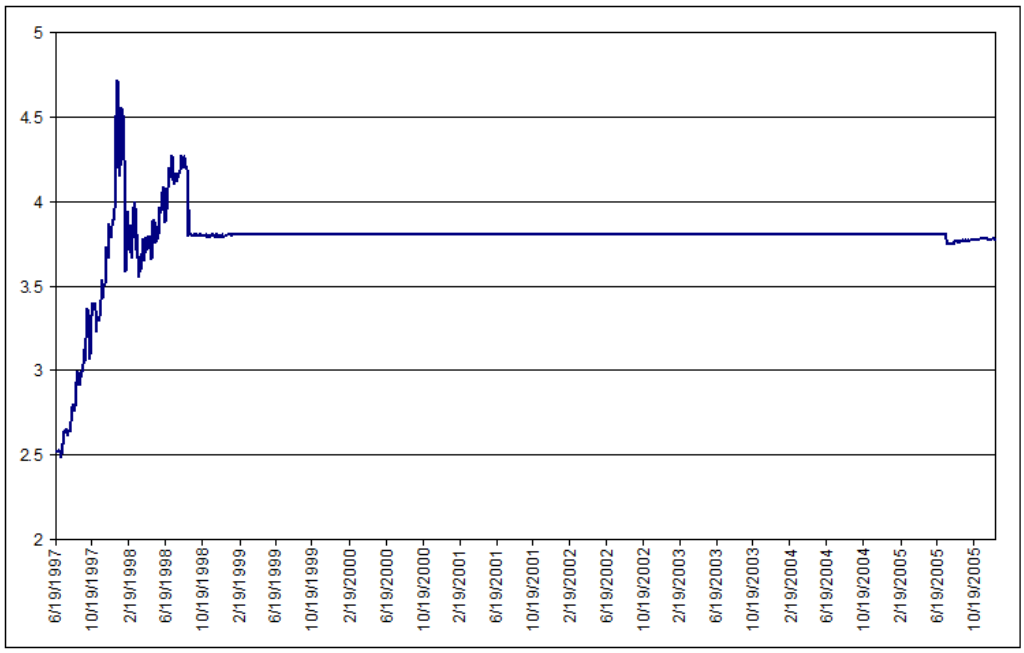
where $\Delta\hat{\theta}_t$ is the change in the vector of regression coefficients resulting from dropping oberservation t, $\hat{\Sigma}$ is the estimated covariance matrix of the coefficients, and k is the number of regressors (including the constant) of the regression. The outliers for regression 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. ***, **, and * indicate significance level at 1%, 5%, and 10%, respectively.

5 Malaysia: The other side of the trinity

In this section, we further illustrate the power of our identification through Impossible Trinity by providing an example at the other side of the trinity. In September 30, 1998, Malaysian government

has responded to the Asian financial crisis in an unorthodox way, compared to other East Asian countries. As shown in Figure 4, the Malaysian Ringgit has been pegged at 3.80 ringgit to the US dollar, but foreign capital repatriated before staying at least twelve months has become subject to substantial levies, and several limitations have been imposed on bank and foreign transactions. The peg has lasted until July 21, 2005, after which Malaysia switched to a managed float against an undisclosed basket of currencies while the capital controls are still in place.[26]

Figure 4: MYR/USD exchange rate



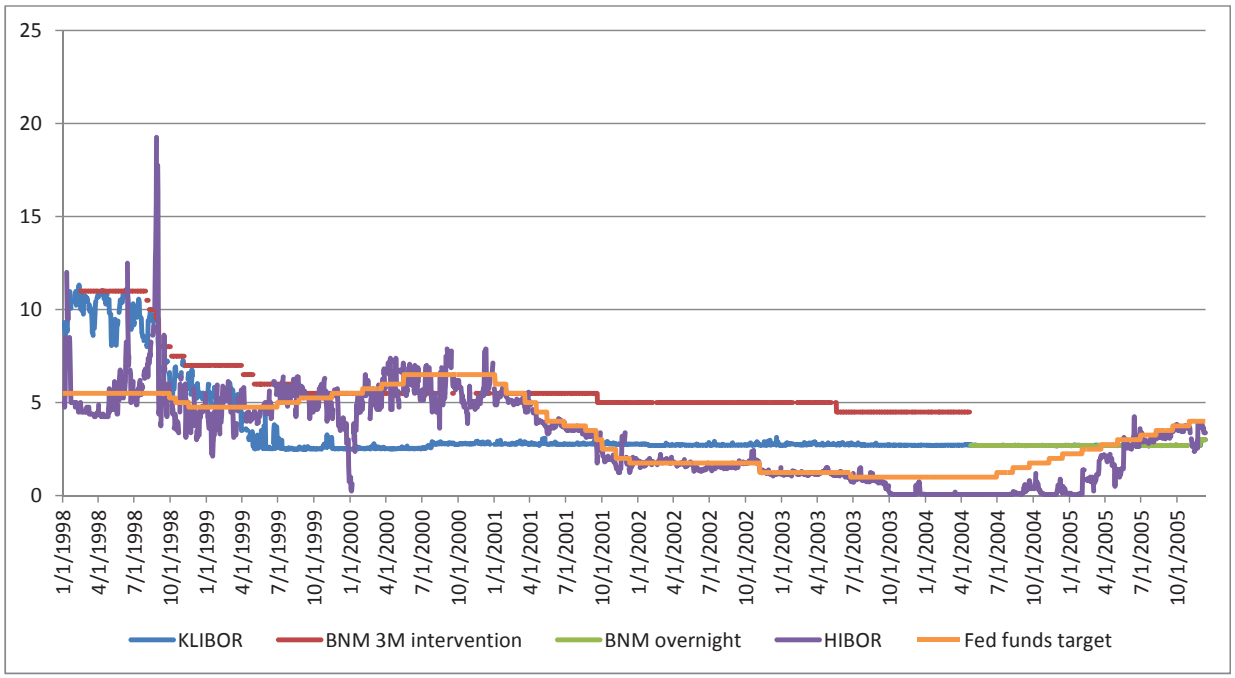
Source: Bloomberg

According to Impossible Trinity theory, the combination of fixed exchange rate and capital controls implies that the monetary policy of Malaysia is independent of the US monetary policy.[27]

[26] 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 in this section. A detailed history of Malaysian capital controls can be found in Johnson, Kochhar, Mitton, and Tamirisa (2007, NBER book).

[27] 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 US and their stock market indices have similar composition. For example, according to IMF DOTS/WEO 2004 data, the ratio of the total trade with US to

Figure 5: Overnight interest rates of Malaysia and Hong Kong



Source: Bloomberg

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 federal funds target rate.[28] Therefore, if our identification strategy has merit we should expect that the stock prices in Malaysia to hardly respond to monetary policy shocks in the US. Table 4 compares the response of the Hang Seng Index and Kuala Lumpur Composite Index to changes in U.S. federal funds target rate. This table clearly fulfills our expectations and hence supports our identification mechanism.

local GDP is 35% for Hong Kong and 32% for Malaysia. Also, the largest 30 firms that constitute more than 80% of the value in Hong Kong and Malaysian stock indices include similar number of banks and financial institutions, 6 in Malaysia and 8 in Hong Kong as of 2004. This makes Malaysia a suitable alternative to study.

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

Table 4: The response of Hong Kong (HSI) and Malaysian (KLCI) stock prices to U.S. federal funds rate changes. The data spans from September 30, 1998 to July 21, 2005, the fixed exchange rate period for Malaysia. The last day of 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.

	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: 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. For the sake of comparability, the outliers are taken as the union of the set of outliers for 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. ***, **, and * indicate significance level at 1%, 5%, and 10%, respectively.

6 Financial Frictions and Monetary Policy Shocks

6.1 A New Testable Implication of Financial Accelerator Theory

We start by showing that the responsiveness of a firm's market value of equity to monetary policy shocks increases as financial frictions increase. We follow the popular framework in Bernanke,

Gertler, Gilchrist (1999), Appendix A in particular. The only difference is that we normalize price of capital and aggregate return on capital to one, since these variables are the same for all firms and we are interested in cross-sectional comparison.

If we let w be the firm's profitability, K be its capital and B be the face value of debt, we can write the firm's problem subject to costly state verification as

$$V = \max_{K,B} E\left(wK - B\right)^{+}$$

subject to the incentive compatibility constraint of the lender

$$R\left(K - N\right) = E\left(\mathbb{I}_{wK \geq B} B + \mathbb{I}_{wK < B}\left(1 - \mu\right) wK\right)$$

where R is the risk-free rate, N is given net worth, or book equity, of the firm, μ is the monitoring cost, and $\mathbb{I}$ denotes the indicator function that is equal to one if the corresponding condition is satisfied and zero otherwise. We are interested in $\partial \ln V / \partial R \partial \mu$.

Defining $v \equiv V/N, k \equiv K/N$, and $\bar{w} \equiv B/K$ we can rewrite the firm's problem as

$$v = \max_{k,\bar{w}} E\left(w - \bar{w}\right)^{+} k$$

subject to

$$R\left(k - 1\right) = E\left(\mathbb{I}_{w \geq \bar{w}} \bar{w} + \mathbb{I}_{w < \bar{w}}\left(1 - \mu\right) w\right) k.$$

We are interested in how the percentage change in stock prices in response to a change in risk-free rate varies with monitoring costs, i.e. we need to find the sign of $\partial \ln v / \partial R \partial \mu$ because net worth, N, is a state variable independent of interest rate. The following proposition shows that under the regularity assumptions in Bernanke, Gertler, Gilchrist (1999) the market values of financially constrained firms are more responsive to interest rate shocks.

Proposition 1 *Let $f(w)$ and $F(w)$ be the pdf and cdf of the firm's productivity, and $h(w) \equiv f(w)/(1-F(w))$ be the hazard rate. The elasticity of market value of equity with respect to risk-free rate is increasing in monitoring cost if $\bar{w}h(\bar{w})$ is increasing in $\bar{w}$.*

Proof. The assumption regarding the hazard rate is imposed by Bernanke, Gertler and Gilchrist (1999) to guarantee a non-rationing outcome. It also guarantees a negative relation of investment to interest rate, which is one of the pillars of financial accelerator theory. We refer the reader to Appendix A.1 of Bernanke, Gertler and Gilchrist (1999) for details. By substituting the incentive compatibility constraint of the lender into the objective function of the firm, we obtain

$$v = \max_{\bar{w}} \frac{R\int_{\bar{w}}^{\infty}(w-\bar{w})\,dF(w)}{R-\left[\bar{w}+\int_0^{\bar{w}}((1-\mu)w-\bar{w})\,dF(\bar{w})\right]} = \frac{RP(\bar{w})}{R-Q(\bar{w},\mu)}.$$

Using first the envelope theorem to find the derivative with respect to R and then by direct differentiation with respect to μ leads to

$$\begin{aligned}\operatorname{sgn}\frac{\partial \ln v}{\partial R\partial\mu} &= -\operatorname{sgn}\frac{dQ(\bar{w},\mu)}{d\mu}\\ &= -\operatorname{sgn}\left(\frac{\partial Q(\bar{w},\mu)}{\partial\bar{w}}\frac{d\bar{w}}{d\mu}+\frac{\partial Q(\bar{w},\mu)}{\partial\mu}\right).\end{aligned}$$

Direct differentiaton shows that $\partial Q/\partial\mu < 0$. Moreover, we have

$$\begin{aligned}\frac{\partial Q}{\partial\bar{w}} &= 1-F(\bar{w})-\mu\bar{w}f(\bar{w})\\ &= [1-F(\bar{w})][1-\mu\bar{w}h(\bar{w})].\end{aligned}$$

Bernanke, Gertler, Gilchrist (1999) show that this expression should be positive in equilibrium if $\bar{w}h(\bar{w})$ is increasing in $\bar{w}$.[29] Finally, we can show $d\bar{w}/d\mu < 0$ by total differentiation of the first

[29] Because $\bar{w}h(\bar{w})$ is increasing in $\bar{w}$ there exists a $\bar{w}^*$ so that $\partial Q/\partial\bar{w} \lesseqqgtr 0$ if $\bar{w} \gtreqqless \bar{w}^*$. Appendix A.1 of Bernanke, Gertler, Gilchrist (1999) shows that $\bar{w} > \bar{w}^*$ cannot be an equilibrium.

order condition of the maximization problem and then making use of its second order condition. Combining these results we have $\partial \ln v / \partial R \partial \mu > 0$ which completes the proof. ■

6.2 Empirical Analysis

There is a large literature that argues cross-listing of foreign firms in U.S. enhances transparency of firms and investor protection, and hence reduces the agency costs.[30] Accordingly, we conclude that Hong Kong firms that are also listed in US exchanges are easier to monitor by the lenders active in US (which forms a much larger market for credit).[31] Moreover, these firms have to satisfy SEC and U.S. exchange listing requirements, including U.S. accounting standards, which provides more information about these firms and makes monitoring easier. Therefore, comparing the cross-listed firms with other firms provides a testing ground for proposition 1. In particular, cross-listed firms should be less financially constrained and hence their prices should be less responsive to monetary policy shocks.

For our analysis, we form a portfolio that starts on July 26, 1993 which is the first cross-listing date we identify. The full list of cross-listing dates is provided in the appendix, Table 9. The value-weighted return of day t, $R(t)$, of this portfolio is given by

$$R(t) = \frac{\sum_i w_i(t-1) r_i(t)}{\sum_i w_i(t-1)},$$

where $r_i(t)$ is the return on stock i on day t and $w_i(t-1)$ is the market capitalization of stock i at the end of day $t-1$. We compare the response of this portfolio to monetary policy shocks with the response of Hang Seng Index (HSI).[32] For the sake of comparability, we only include the event

[30] A non-exhaustive list includes Coffee (1999, 2002), Stulz (1999), Reese and Weisbach (2002), and Doidge, Karolyi, and Stulz (2004).

[31] The monitoring benefit of cross-listing, also called "bonding", exists even in perfectly integrated capital markets. The key for this benefit to exist is that investors in a firm become better protected if the firm lists in the U.S. This can give lenders an edge in the bankruptcy courts, especially given that Hong Kong has still not adopted a procedure similar to Chapter 15 in the US that deals with the resolution of cross-border insolvency cases.

[32] Daily data on return and capitalization can be accessed through Bloomberg. We make sure that stocks are not part of both HSI index and our cross-listed portfolio at the same time. So, HSI* in Table 5 is a value-weighted portfolio of

dates for which both returns are available. Table 5 reports the results which do not seem to support proposition 1 because cross-listed firms are not less responsive than single-listed firms.[33]

As an additional test of proposition 1, we consider firm size as a proxy for financial constraints, following Gertler and Gilchrist (1994) who use this proxy to show that financial frictions amplify the response of investment to a tightening of monetary policy. We use MSCI Hong Kong Small Cap Index (MXHKSC) as the portfolio of small firms and compare it with HSI index which includes larger firms.[34] According to proposition 1, we should expect that the small firms are more responsive to the monetary policy shock. However, we do not see this result in Table 6 which shows that the index of small stocks is less responsive than the HSI index. One possible reason for this phenomenon might be that small firms' stock prices respond with a delay because they are less frequently traded. To control for this possibility, we replace our daily return variables with weekly returns and present the results in Table 10 of the appendix. While the response of small firms' prices seems to catch up with the response of large firms' prices over the weekly period, their response is never greater. We obtain qualitatively similar results when we extend the return horizon to two weeks or a month.

To summarize, we cannot find evidence in favor of proposition 1, which contradicts the financial accelerator channel of monetary policy.

the HSI constituents that are not cross-listed.

[33]One could argue it is not surprising that cross-listed firms are more responsive to U.S. monetary policy shocks since their fate might be more closely tied to U.S. economy. However, Table 5 also indicates that the stock prices of the cross-listed firms are less responsive to movements in U.S. equity market (0.38 versus 0.54), providing evidence against this claim.

[34]The data for MXHKSC comes from Bloomberg and has the starting date February 1995.

Table 5: Cross-listed firm portfolio (1993-2008) if the stock is in cross-listed, it is taken out of HSI

	Full sample		Excluding outliers	
Regressor	portfolio	HSI*	portfolio	HSI*
Intercept	0.23	0.21	0.16	0.16
	(0.21)	(0.14)	(0.21)	(0.13)
Expected change	1.00	0.60	1.44	0.30
	(1.01)	(0.78)	(1.08)	(0.68)
Surprise change	-8.37***	-7.51***	-9.18***	-2.75*
	(2.33)	(2.85)	(3.15)	(1.55)
US equity returns	0.38**	0.65***	0.36*	0.54***
	(0.16)	(0.15)	(0.19)	(0.12)
Adjusted R^2	0.17	0.36	0.07	0.14
Obs.	113	113	107	107
χ^2: porfolio = HSI	0.26		5.14**	

Note: HSI* is a value-weighted portfolio of the HSI constituents that are not cross-listed. 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 the HSI regression are May 17, 1994, October 15, 1998, January 3, 2001, April 18, 2001, January 22, 2008, and March 18, 2008. For the sake of comparability, the outliers for the portfolio regression are the same as those for the HSI regression. The last row of this table report χ^2 obtained from the post-estimation of the seemingly unrelated regression (SUR) system consisting of porfolio and HSI equations. The post-estimation is on the coefficient "Surprise change". Robust standard errors are reported in parentheses. ***, **, and * indicate significance level at 1%, 5%, and 10%, respectively.

Table 6: The response of small firms index versus aggregate stock indices for Hong Kong (1995-2008)

	Full sample		Excluding outliers	
Regressor	MXHKSC	HSI	MXHKSC	HSI
Intercept	0.17	0.07	0.19*	0.08
	(0.11)	(0.12)	(0.11)	(0.12)
Expected change	-0.50	-0.48	-0.46	-0.47
	(0.51)	(0.62)	(0.58)	(0.64)
Surprise change	-4.28***	-10.32***	-1.10	-5.29***
	(0.85)	(2.51)	(1.39)	(1.75)
US equity returns	0.45***	0.52***	0.38***	0.52***
	(0.07)	(0.14)	(0.09)	(0.12)
Adjusted R^2	0.35	0.50	0.10	0.17
Obs.	102	102	97	97
χ^2: MXHKSC = HSI	31.05***		4.23**	

Note: 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 the HSI regression are October 15, 1998, January 3, 2001, April 18, 2001, January 22, 2008, and March 18, 2008. For the sake of comparability, the outliers for the MXHKSC regression are the same as those for the HSI regression. The last row of this table report χ^2 obtained from the post-estimation of the seemingly unrelated regression (SUR) system consisting of the MXHKSC and HSI equations. The post-estimation is on the coefficient "Surprise change". ***, **, and * indicate significance level at 1%, 5%, and 10%, respectively.

7 Conclusion

On the basis of 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 US federal funds target rate. As summarized in Rigobon and Sack (2004), two major problems arise in estimating stock market's response to monetary policy. One is that monetary policy is simultaneously influenced by fluctuations in stock market. The other is that there may be factors that have a direct impact on both monetary policy and stock market, which creates an omitted variable bias. By focusing on Hong Kong stock market's response to US monetary policy, we circumvent the simultaneity problem, since changes in Hong Kong stock prices do not directly influence US monetary policy. We also show that using US stock returns as an additional regressor controls for omitted variable bias.

In addition, we reveal and test a new implication of broad credit channel of monetary policy: The stock prices of firms with higher external finance cost are more responsive to monetary policy shocks. This implication of credit channel is not supported by our analysis which relies on comparing firms that are cross-listed in US and Hong Kong with the firms only listed in Hong Kong. We also use firm size as a proxy for the degree of financial constraint as in Gertler and Gilchrist (1994) and reach a similar result. We obtain similar results (not reported here) when we use the size portfolios from Ken French's website for the NYSE-AMEX-NASDAQ universe. This finding is also consistent with Lamont, Polk, and Saá-Requejo (2001) who find that the relative performance of constrained firms does not reflect monetary policy or credit conditions.

These results are in contrast with the broad implications of the financial accelerator model of Bernanke, Gertler, and Gilchrist (1999) and warrants further analysis. Our next step is comparing firms with and without bond ratings as in Kashyap, Lamont, and Stein (1994) and using alternative financial frictions indices such as Kaplan and Zingales (1997) or Whited and Wu (2006).

8 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

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

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 observation. Then, the OLS estimates for the parameters

a, b, c are given by

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

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, this restriction does not have any economic justification.[35] 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), we do not need homoscedasticity of non-monetary shocks and variables for our identification mechanism.

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 $e = 0$.

[35] The exclusion of non-linear 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 $1 - \alpha\beta \neq 0$.

8.1 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

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

where $\sigma_{i,t} = var\left(\Delta i_t\right), \sigma_{s,t} = var\left(\Delta s_t\right), \sigma_{i,s,t} = cov\left(\Delta i_t, \Delta s_t\right), \sigma_{i,ez,t} = cov\left(\Delta i_t, ez_t\right)$ and $\sigma_{s,ez,t} = cov\left(\Delta s_t, ez_t\right)$. 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

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

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

The results, shown in Table 7, 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 have not used the pre-event dates. We do not observe such a difference between the two estimators which is also consistent with our hypothesis that $e = 0$.

Table 7: The response of HK equity returns to US federal funds rate changes and US equity returns (1994-2005). The policy shocks on pre-event dates are calcuated in the same way as Kuttner's surprises.

	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: 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. ***, **, and * indicate significance level at 1%, 5%, and 10%, respectively. The data to calculate the policy surprises at pre-event dates comes from Bloomberg, with the exception of 14 November 1994 and 19 December 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 8: 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 fed funds rate. 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. 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. ***, **, and * indicate significance level at 1%, 5%, and 10%, respectively.

Table 9: HK-US cross-listed companies and their U.S. exchange and HSI listing dates

Company name	HK symbol	US symbol	U.S. Listing	HSI Listing
China Eastern Airlines Corp. Ltd.	0670 HK	CEA	2/4/1997	
CNOOC Ltd.	0883 HK	CEO	2/27/2001	7/31/2001
City Telecom HK Ltd.	1137HK	CTEL	11/3/1999	
Yanzhou Coal Mining Co. Ltd.	1171HK	YZC	3/31/1998	
China Petroleum & Chemical Corp.	0386HK	SNP	10/18/2000	12/4/2006
PetroChina Co. Ltd.	0857HK	PTR	4/4/2000	12/10/2007
Huaneng Power Intl. Inc.	0902HK	HNP	10/ 6/1994	
Aluminum Corporation of China Limited	2600HK	ACH	12/11/2001	6/10/2008
HSBC Holdings plc	0005HK	HBC	7/16/1999	4/3/1991
China Life Insurance Co. Ltd.	2628HK	LFC	12/17/2003	3/12/2007
Semiconductor Manufacturing Intl.Corp.	0981HK	SMI	3/17/2004	
China Southern Airlines Co. Ltd.	1055HK	ZNH	7/30/1997	
China Unicom (Hong Kong) Limited	0762HK	CHU	6/21/2000	6/1/2001
China Telecom Corp. Ltd.	0728HK	CHA	11/14/2002	
Guangshen Railway Co. Ltd.	0525HK	GSH	5/13/1996	
Sinopec Shanghai Petrochemical Co. Ltd.	0338HK	SHI	7/26/1993	
China Mobile Limited	0941HK	CHL	10/22/1997	1/27/1998

Table 10: Size (1995-2008) 7 day with interpolated CRSP return

	Full sample		Excluding outliers	
Regressor	MXHKSC	HSI	MXHKSC	HSI
Intercept	0.22	-0.06	0.15	-0.10
	(0.28)	(0.24)	(0.29)	(0.24)
Expected change	0.66	0.58	1.15	1.20
	(0.96)	(0.94)	(1.22)	(0.99)
Surprise change	-7.30***	-9.76***	-11.84***	-11.98***
	(2.10)	(1.82)	(2.79)	(2.27)
US equity returns	0.74***	0.95***	0.74***	0.92***
	(0.12)	(0.09)	(0.12)	(0.09)
Adjusted R^2	0.35	0.54	0.34	0.50
Obs.	102	102	99	99
χ^2: MXHKSC = HSI	1.19		0.00	

Note: 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 the HSI regression are April 18, 2001, January 22, 2008, and March 18, 2008. For the sake of comparability, the outliers for the MXHKSC regressions are the same as those for the HSI regression. The last row of this table report χ^2 obtained from the post-estimation of the seemingly unrelated regression (SUR) system consisting of the MXHKSC and HSI equations. The post-estimation is on the coefficient "Surprise change". ***, **, and * indicate significance level at 1%, 5%, and 10%, respectively.

9 References

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