THREE ESSAYS ON THE GLOBAL INFLUENCE OF U.S. MONETARY POLICY

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Kim, Kyoung-Gon (Ph.D., Economics)

Three Essays on the Global Influence of U.S. Monetary Policy

Thesis directed by Professor Martin Boileau

In the first chapter, I identify how the Fed's dependence on unconventional monetary policy after the 2007-2008 financial crisis and its return to conventional policy in 2015 have affected the global influence of U.S. monetary policy. I divide the sample into three phases according to the Fed's monetary policy regimes: pre-crisis (Aug 2001 - Nov 2008), crisis (Nov 2008 - Dec 2015), and post-crisis (Dec 2015 - Sep 2017). Daily variations in government bond yields and foreign exchange spot rates for 46 countries on FOMC meeting days show that the influence of U.S. monetary policy surprises intensified after the financial crisis. Responses are stronger in a group of emerging markets than in developed economies. I also find that more flexible exchange rate regimes lead to larger magnitudes of responses to U.S. monetary policy surprises. My results show that the decoupling of interest rates between the U.S. and other countries forced foreign financial markets to respond sensitively to U.S. monetary policy surprises after the financial crisis.

In the second chapter, I examine whether the global transmission of U.S. monetary policy surprises to stock price indexes and term spreads in G7 economies changed after the 2007-2008 financial crisis. I estimate a vector error correction model using monthly data spanning 2001-2017. I find that monetary tightening induces a reduction in stock price indexes and term spreads before the crisis. This confirms the conventional view of the effects of monetary policy on stock and bond markets. However, an unanticipated tightening in U.S. monetary policy leads to an increase in stock price indexes and term spreads in the post-crisis period. This positive response is at odds with the conventional view. A plausible explanation attributes a role to a bubble component of asset prices. Keeping interest rates close to the zero lower bound for many years in G7 countries may have led to a lower borrowing cost, which would presumably increase the size of an asset bubble. As a result, the Fed's tapering of quantitative easing and raising the Fed Fund rates since 2015 would lead to a surge in stock prices.

In the final chapter, I investigate the effect of real exchange rates on international trade through monetary policy. I estimate a vector autoregression model using monthly data from China, Japan, and Korea spanning 2001-2017. I find that an innovation in U.S. monetary policy shocks leads to an immediate increase in the trade volume in Korea. However, real effective exchange rates and trade volumes move to same direction in China, which is at odds with the conventional view on the relationship between exchange rate and trade. Likewise,

a depreciation of local currency due to a contractionary U.S. monetary policy improves the trade balance in Korea but it leads to a deterioration in the trade balance in China. The heterogeneous responses in Korea and China may be attributed to the different extent of global value chain participation between large and small open economies.

Dedicated to my loving wife and best friend, Dure

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Chapter 1

Financial Crisis and the Global Transmission of U.S. Monetary Policy Surprises

1.1 Introduction

As global capital markets integrate, U.S. monetary policy is more likely to affect the economies of other countries. As a result, finance ministers in emerging markets often worry that their economies are influenced by U.S. monetary policy. For example, the Federal Reserve's reduction of the Fed Funds rate to its effective lower bound on December 16, 2008, led to a decrease in government bond yields and an appreciation of local currencies in more than 30 countries for one day. This is why global financial markets pay attention to Fed announcements on the day the Federal Open Market Committee (FOMC) meets. In this paper, I investigate whether the 2007-2008 U.S. financial crisis changed the influence of the Fed's surprising decisions on foreign financial markets. Specifically, I focus on how the Fed's dependence on unconventional monetary policy after the financial crisis and its return to conventional policy in 2015 affected the global influence of U.S. monetary policy surprises. Using daily variations in government bond yields and foreign exchange spot rates for 46 sample countries on FOMC meeting days, I find that the global influence of U.S.

monetary policy surprises intensified after the financial crisis: The widening gap in interest rates between the U.S. and the rest of the world rendered foreign financial markets more sensitive to Fed decisions after the crisis.

The financial crisis led to a global economic downturn and a European debt crisis. The Fed responded aggressively to the crisis by lowering the Fed Funds rate to a range between 0 and 0.25 percent, the lowest in its history. Also, the Fed adopted unconventional policies: forward guidance on future interest rates and quantitative easing (QE) with large-scale asset purchases (LSAP). The Fed eventually escaped the zero lower bound (ZLB) in December 2015 by raising the Fed Funds rate for the first time since 2006. As of November 2017, the Fed has raised the target range for the Fed Funds rate to between 1.00 and 1.25 percent. I design an empirical model that employs data covering all FOMC meetings from August 2001 to September 2017. I divide this sample into three phases according to Fed monetary policy regimes: pre-crisis, crisis, and post-crisis. I assume that unconventional monetary policies due to the financial crisis began when the Fed's plan for large-scale asset purchases (LSAP-I) was announced (November 25, 2008) and ended when the Fed raised the Fed Funds rate again (December 16, 2015).[1]

I calculate U.S. monetary policy surprises by changes in the response of U.S. financial markets to the Fed's decision. To do so, I use high-frequency tick data for two types of futures, Fed Funds futures and 10-year Treasury futures, around the announcement of the Fed's decision (2:15 pm ET). Fed Funds futures are financial contracts that reflect market views on the likelihood of Fed policy changes. These contracts have a payout based on the average effective Fed Funds rate that prevails over the calendar month specified in the contract. I define a *Fed Funds futures surprise* by the changes in the Fed Funds futures rate between 10 minutes before and 20 minutes after an FOMC announcement. Within this 30-minute window, the Fed Funds futures surprise measures the unanticipated component of

[1] Gilchrist et al. (2015b) regard November 25, 2008, as the key date on which the Fed announced its plan for buying the debt obligations of government-sponsored enterprises (GSEs) and mortgage-backed securities (MBS) for the first time. In this study I follow their assumption that the unconventional monetary policy began on November 25, 2008.

the Fed's decision on the Fed Funds rate target (Kuttner (2001)). Ten-year Treasury futures are derivatives whose prices are closely tied to the prices of U.S. Ten-year government bonds and their yields. Ten-year Treasury bonds carry almost zero risk to the principal, and are thus considered to be an important measuring stick for market confidence about the future. I calculate a *Treasury futures surprise* by changes in the 10-year Treasury futures price within a 30-minute window around a FOMC announcement. The Treasury futures surprise captures the future path of expected interest rates contained in the Fed's announcement.

I measure the responses of foreign financial markets to U.S. monetary policy surprises by daily variations in government bond yields and foreign exchange spot rates in 46 countries on FOMC days. I examine how short-term (2-year), midterm (5-year), and long-term (10-year) sovereign bond yields respond to U.S. monetary shocks in pre-crisis, crisis, and post-crisis periods. My estimates indicate that the response of sovereign bond yields to U.S. monetary policy surprises differs not only across maturities, but also across periods. For an unanticipated increase in Fed Funds futures by 100 basis points, the yields on long-term sovereign bonds in the post-crisis period rise by an additional 120 basis points, relative to bond yields in the pre-crisis period. Likewise, an unexpected decrease in Fed Funds futures by 100 basis points leads to a decline in the yield on short-term sovereign bonds by 80 additional basis points in the crisis period compared to the pre-crisis period.

Next, I investigate the relationship between foreign exchange spot rates and U.S. monetary policy surprises. My estimates show that a decline in the Fed Funds futures surprise of 100 basis points is associated with an appreciation in the local currencies of an additional 9 percent in the crisis period and an extra 16 percent in the post-crisis period compared to the pre-crisis period. I attribute this to the decoupling of interest rates between the U.S. and other countries. In the face of the financial crisis, the Fed cooperated with other central banks to prevent a deepening of the global credit crisis. However, when the Fed raised the Fed Funds rate in 2015, the policy coordination cracked; Europe and Japan kept their rates near zero. Central banks in emerging markets also didn't pursue premature monetary tightening. The widening interest rate gap between the U.S. and the rest of the world forced foreign financial markets to respond sensitively to the Fed's decision.

In an effort to identify whether emerging markets are more vulnerable to U.S. monetary policy shocks, I divide the sample of countries into two groups: developed economies and emerging markets. Overall estimates indicate that responses to U.S. monetary policy surprises are stronger in emerging markets than in developed economies. This finding is consistent with those reported by Chen et al. (2016). When taking into account exchange rate regimes (hard pegs, soft pegs, managed float, and free float), I find that free-floating arrangements lead to the larger responses to U.S. monetary policy surprises. Under a free-floating regime, a rise in the Fed Funds futures surprise by 100 basis points leads to an increase in the 10-year government bond yield of 20 additional basis points in the crisis period and 68 extra basis points in the post-crisis period compared to the pre-crisis period. However, the results present no significant response of government bond yields under hard-pegged regimes.

My findings are robust to various additional tests. First, I address the possible non-independence of error terms by clustering standard errors. While government bond yields and foreign exchange rates change at a country level, U.S. monetary policy surprises vary at an aggregate level in my data. This may lead to the nonindependence of error terms for each FOMC meeting, which would underestimate standard errors. Clustering at the FOMC meeting level confirms that the global influence of U.S. monetary policy surprises intensified after the financial crisis. Second, I isolate the component of changes in the 10-year Treasury futures price that is not related to the Fed Funds futures surprise. I define the *Residual surprise* as the error term from the regression of Treasury futures surprise on the Fed Funds futures surprise. The Residual surprise reflects the expected future path of interest contained in the FOMC announcement that is orthogonal to the movement in Fed Funds futures (Wongswan (2009)). A bootstrapped two-step estimation method suggests that the responses of government bond yields and exchange rates to Fed Funds futures and Residual surprises become stronger after the financial crisis.

This paper contributes to the empirical literature that explores the global spillovers of U.S. monetary policy in several ways. The first contribution is showing that the financial crisis affected the influence of U.S. monetary policy surprises. The Fed's dependence on QE in the financial crisis led to a voluminous literature on how unconventional U.S. monetary policy affects global economies (Banerjee et al. (2016); Gilchrist et al. (2015b); Chen et al. (2016); Gagnon et al. (2017); Meinusch and Tillmann (2016); Bowman et al. (2015); Bauer and Neely (2014); Neely (2015); Swanson and Williams (2014)). Banerjee et al. (2016) show that unexpected U.S. monetary policy tightening leads to a fall in GDP, rise in interest rates, and depreciation in exchange rates in emerging market economies. Meinusch and Tillmann (2016) empirically find that QE is associated with higher output and inflation and lower nominal interest rates in U.S. However, Gagnon et al. (2017) find that U.S. unconventional monetary policy weakens the connection between U.S. bond yields and foreign currencies. To my knowledge, my paper is the first to identify different responses to U.S. monetary policy surprises, not only during the crisis but also in the post-crisis period, using high-frequency data.

The second contribution is highlighting the relationship between U.S. monetary policy and exchange rate regimes. Aizenman et al. (2017) show that the type of exchange rate regime matters for the transmission of shocks. Hausman and Wongswan (2011) show that interest rates in less flexible regimes respond more to U.S. monetary policy. Bowman et al. (2015) find that sovereign bond yields in a managed floating currency are more exposed to changes in U.S. financial conditions than those in free-floating currencies. I show empirically that within a 1-day window, the responses of exchange rates and sovereign bond yields to U.S. monetary shocks are greater under the free-floating exchange rate regime than those in fixed exchange rate regimes.

The third contribution is showing that the magnitude of spillovers is different for developed economies and emerging markets. Gilchrist et al. (2015a) find that U.S. monetary policy has a bigger effect on short- and long-term interest rates for developed economies relative to emerging markets. However, Chen et al. (2016) show that emerging markets are more likely to respond to QE when using monthly data between 2007 and 2013. I add empirical evidence that the responses of emerging markets to a U.S. monetary policy surprise became stronger than those of developed economies after the financial crisis.

The remainder of the paper is organized as follows. Section 1.2 discusses the background of the study. Section 1.3 describes the data and methodology. Section 1.4 presents the results for spillover estimates of U.S. monetary policy surprises. Section 1.5 tests the robustness of the results, and Section 1.6 concludes.

1.2 Background

1.2.1 Global Transmission Channels of U.S. Monetary Policy

Uncovered nominal interest parity explains how exchange rates respond to changes in interest rates caused by monetary policy. The theory states that expected changes in the exchange rate depend on interest rate differentials:

$$E_t s_{t+1} - s_t = i_t - i_t^*, \tag{1.1}$$

where s_t is the nominal exchange rate between two currencies at time t, $E_t s_{t+1}$ is an expected value of s_{t+1} with the information available at time t, and i_t is the nominal interest rate in the home country (similarly, i_t^* is for the foreign country).[2] If the home country has a higher nominal interest rate (i.e., $i_t > i_t^*$), its currency is expected to depreciate (i.e., a rise in s) to equalize returns in the two countries. Under rational expectation, the exchange rate at $t+1$ can be expressed as the sum of the expected value of the exchange rate and a forecast error (φ_t):

$$s_{t+1} = E_t s_{t+1} + \varphi_t. \tag{1.2}$$

[2] s_t is measured by the price of foreign currency in terms of domestic currency. A rise in s_t implies depreciation of the domestic currency

Thus, uncovered interest parity (UIP) can be written as

$$s_{t+1} - s_t = a + b(i_t - i_t^*) + \varphi_{t+1}, \tag{1.3}$$

where $a = 0$ and $b = 1$. However, empirical evidence shows that $b < 0$: Currencies with high interest rates will appreciate, not depreciate (Boudoukh et al. (2016)). This suggests that one can profit from using a carry trade. That is, investors borrow in low interest currencies and invest in higher interest currencies.

For example, when the Fed tightens its monetary policy, nominal interest rates in U.S. rise in the short run. According to carry trade activity, carry traders want to buy more U.S. bonds because U.S. bonds pay a higher interest rate than before (Anzuini and Fornari (2012)). As the demand for dollars to buy U.S. bonds increases, the dollar appreciates in the short run.[3] Figures 1.1 and 1.2 indicate that foreign government bond yields and exchange rates respond to the Fed's announcement in the direction forecast by carry trade activity. On December 16, 2008, the Fed decided to lower the Fed Funds rate to the range between zero and 0.25 percent. The decrease in the Fed Funds rate instantly led to a decrease in 2-year government bond yields and appreciation of local currencies in more than 30 countries for one day, as shown in Figure 1.1. After 4.5 years, on June 19, 2013, the Fed announced a "tapering" of quantitative easing (QE) policies by scaling back its bond purchases. On this day, the global financial market interpreted the announcement as a signal that the Fed would raise the Fed Funds rate in the future. As a result, government bond yields increased and local currencies depreciated in 34 countries for one day, as shown in Figure 1.2.

Several other channels may also affect spillover of U.S. monetary policy (Rey (2016); Borio and Zhu (2012)). For example, according to the credit channel, when the Fed relaxes its monetary policy, nominal interest rates drop, and this leads to an increase in the equity price. As a result, the net worth of borrowers rises and global banks' lending increases.

[3]Two main conditions for carry trade are low exchange rate volatility and high interest rate differentials across countries.

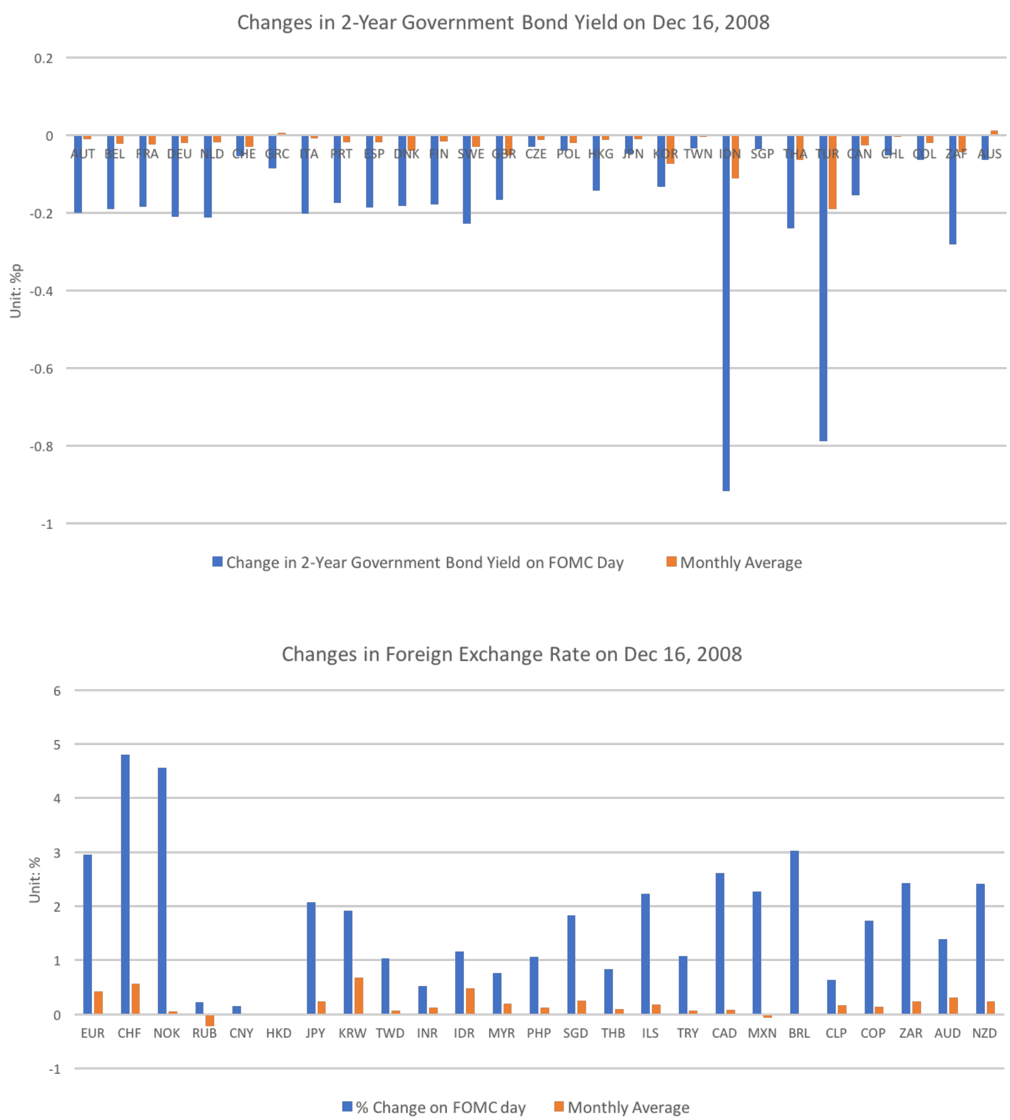


Figure 1.1: Changes in Foreign Government Bond Yields and Exchange Rates on Dec 16, 2008

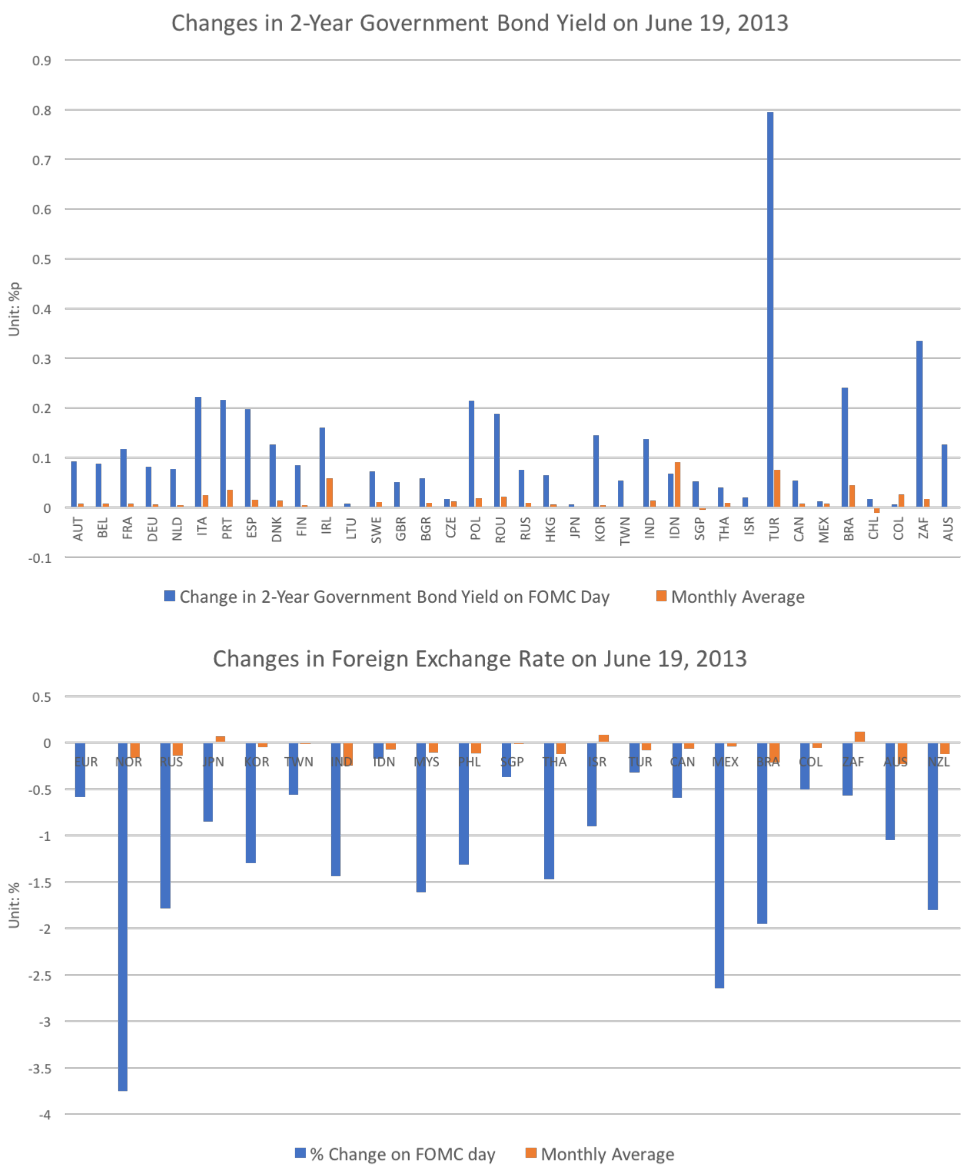


Figure 1.2: Changes in Foreign Government Bond Yields and Exchange Rates on June 19, 2013

This could explain the positive correlation between short-term rates in foreign countries and the Fed Funds rate. The risk-taking channel has a similar mechanism. Relaxation of U.S. monetary policy leads to drops in nominal interest rates. As the returns from safe assets decrease, banks apply relatively low credit standards. Accordingly, the global credit supply goes up and short-term rates in foreign countries move downward. Lastly, the balance sheet channel shows that even advanced economies cannot be free from the influence of U.S. monetary policy. When the Fed tightens its monetary policy, a foreign country's domestic currency depreciates. This helps increase the foreign country's exports. However, as banks become more cautious of the rising (dollar-denominated) value of foreign debt, interest rates rise and bank loans may decrease.

The empirical question is whether we can extend the response of foreign government bond yields and exchange rates to the Fed's decision to all FOMC meetings. If so, how much does U.S. monetary policy influence the movement in foreign government bond yields and exchange rates?

1.2.2 The Financial Crisis and Monetary Policy Regime

As shown in Figure 1.3, the 2007-2008 financial crisis was a huge turning point in the Fed's history. Before the crisis, the Fed managed the Fed Funds rate as a key instrument for its monetary policy. For example, on June 25, 2003, the Fed cut the Fed Funds rate by a 0.25 percentage point to 1 percent, the lowest level in 45 years, to overcome the 2001 recession. The very low interest rates led to a housing boom, solid pace of economic expansion, and improved labor market conditions. As a result, the Fed raised the Fed Funds rate to 1.25 percent on June 30, 2004, which was the first increase since 2000.

However, the 2007-2008 financial crisis, triggered by the bursting of the subprime mortgage bubble and the collapse of Lehman Brothers, dramatically changed the Fed's policy regime, as shown in Table 1.1. On December 16, 2008, the Fed responded aggressively to the crisis by dramatically lowering the Fed funds rate to "between 1/4 points and zero," the

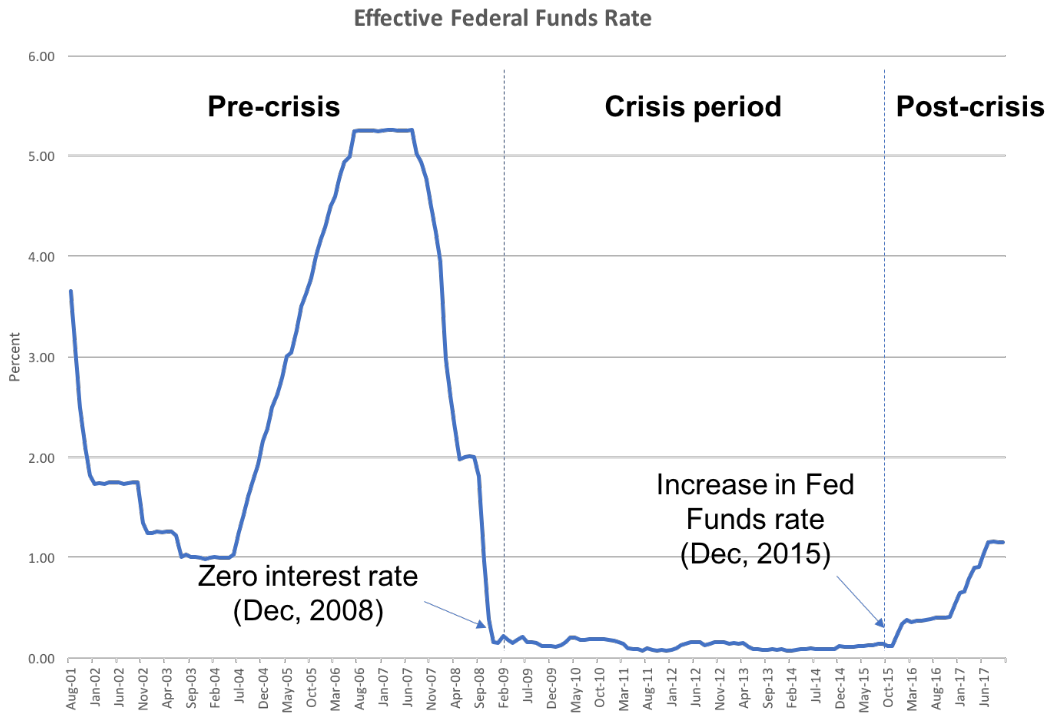


Figure 1.3: Movement of Effective Fed Funds Rate

lowest rate in its history. Facing the ZLB, the Fed had no room for additional moves in the Fed Funds rate if the economy did not improve soon. As a result, instead of adjusting the Fed Funds rate, the Fed adopted unconventional policies, such as forward guidance on future interest rates and QE with LSAP to stimulate the economy and keep market rates low. It tried to influence expectations for the future path of Federal Funds rates through the FOMC statement, a press release, and the chairperson's public speech. The Fed also cooperated with other central banks to prevent further deepening of the global credit crisis. For example, on October 8, 2008, the Federal Reserve and the central banks of the E.U., U.K., Canada, Sweden, and Switzerland cut their rates by one-half point. One week later, the U.S., E.U., and Japan also adopted a coordinated policy to prevent banks from failing. The unconventional monetary policy regime ended in December 2015, when the Fed raised the Fed Funds rate for the first time since 2006. This action officially marks "the end of an extraordinary seven-year period during which the Federal Funds rate was held near zero to support the recovery of the economy from the worst financial crisis and recession since the

Time	Event
Feb 2007	Home sales peak
Mar 2007	Hedge funds housing losses spread subprime misery
Apr 2007	Help for homeowners not enough
Aug 2007	Fed lowers rate to 4.75%
Sep 2007	LIBOR rate unexpectedly diverges
Nov 2007	Treasury creates $75 billion superfund
Dec 2007	Foreclosure rates double
Jan 2008	Fed tries to stop housing bust
Mar 2008	Fed begins bailouts
Apr 2008	Fed lowers rate to 2%
Sep 2008	Lehman Brothers bankruptcy triggers global panic
Sep 2008	Paulson and Bernanke submit bailout to Congress
Oct 2008	Central banks coordinate global action
Nov 2008	Announcement of Large Scale Asset Purchase (LSAP-I)
Dec 2008	Zero interest rates
Nov 2010	Announcement of LSAP-II
Sep 2012	Announcement of LSAP-III
Jun 2013	Announcement of "tapering"
Dec 2015	Fed raises rate for the first time since 2006

Table 1.1: Timeline of the Financial Crisis

Great Depression."[4] Since then, as of November 2017, the Fed has raised the Fed Funds rate three times to the range of 1.00 to 1.25.

The question is how has the Fed's dependence on unconventional monetary policy after the financial crisis, and its return to conventional policy in 2015, affected the global influence of U.S. monetary policy? To address this question, I divide the sample into three phases: pre-crisis, crisis, and post-crisis. I assume that the financial crisis period began when the Fed's LSAP-I plan was announced (November 25, 2008) and ended when the Fed raised the Fed Funds rate again (December 16, 2015).

[4]Transcript of Fed Chair Janet Yellen's press conference, December 16, 2015.

1.3 Empirical Analysis

1.3.1 Monetary Policy Surprises

I measure U.S. monetary policy surprises by changes in the response of U.S. financial markets to the Fed's decision. For this, I collect high-frequency tick data for two types of futures: Fed Funds futures and 10-year Treasury futures.

Fed Funds futures are financial contracts that reflect market views of the likelihood of Fed policy changes. The contracts have a payout based on the average effective Fed Funds rate that prevails over the calendar month specified in the contract. The Fed Funds futures rate 10 minutes before ($f_{t,-10}$) the FOMC announcement (2:15 pm, ET) on day d of a month with D days is calculated by the *average* of the effective overnight Fed Funds rate as follows:

$$f_{t,-10} = \frac{d(Realized) + (D-d)(Expected_{t,-10})}{D}, \tag{1.4}$$

where $Realized$ is the effective Fed Funds rates during the past d days of the relevant month and $Expected_{t,-10}$ is the expectation of the Fed Funds rate for upcoming $D-d$ days of the month 10 minutes before the FOMC announcement. In Equation (1.4), I solve for $Expected_{t,-10}$ to factor out the market's expectation for the Fed's decision before the announcement:

$$Expected_{t,-10} = \frac{D}{D-d}(f_{t,-10}) - \frac{d}{D-d}(Realized). \tag{1.5}$$

Similarly, I calculate the expected value $Expected_{t,+20}$ for the Fed Funds rate for forthcoming $D-d$ days of the month 20 minutes after the FOMC announcement:

$$Expected_{t,+20} = \frac{D}{D-d}(f_{t,+20}) - \frac{d}{D-d}(Realized), \tag{1.6}$$

where $f_{t,+20}$ (the Fed Funds future rate 20 minutes after the FOMC announcement) reflects how the financial markets interpreted the Fed's decision ex post.

I define a Fed Funds futures surprise, FF_t, by changes in the expectation for the Fed Funds rate between 10 minutes before ($Expected_{t,-10}$) and 20 minutes after ($Expected_{t,+20}$) the FOMC announcement from Equations (1.5) and (1.6):

$$FF_t = \frac{D}{D-d}(f_{t,+20} - f_{t,-10}). \tag{1.7}$$

Within a 30-minute window, the Fed Funds futures surprise (FF_t) measures the unanticipated component of the Fed's decision on the current Fed Funds rate target (Kuttner (2001); Gertler and Karadi (2015)). If there is no surprise in the Fed's decision, FF_t is zero, because $f_{t,-10}$ and $f_{t,+20}$ have the same value.

However, when the Fed Funds rate dropped to its ZLB in the financial crisis period, changes in the current Fed Funds future rate might be restricted. To address this problem, I employ 10-year Treasury futures that reflect a future path for monetary policy contained in the FOMC statement. Ten-year Treasury futures are derivatives whose prices are closely tied to the prices of U.S. 10-year government bonds and their yields. Ten-year Treasury bonds carry almost zero risk to principal, and thus, are considered to be an important measuring stick for market confidence about the future. For example, when confidence is high, the 10-year Treasury bond's price drops and yields go higher. I calculate a Treasury futures surprise, TYF_t, by changes in the 10-year Treasury futures price between 10 minutes before ($tyf_{t,-10}$) and 20 minutes after ($tyf_{t,+20}$) the FOMC announcement, as follows:

$$TYF_t = tyf_{t,+20} - tyf_{t,-10}. \tag{1.8}$$

Gürkaynak et al. (2005) find that 75 to 90 percent of variations in 10-year Treasury yields respond to forward guidance in FOMC statements rather than the current Fed Funds rate target. Therefore, changes in the 10-year Treasury futures price within a 30-minute window around an FOMC announcement (TYF_t) capture the future path of expected interest rates contained in FOMC statements.

The sample period in my dataset includes all FOMC meetings from August 2001 to September 2017. The FOMC holds eight regularly scheduled meetings each year. In addition, the FOMC holds irregular intermeetings as needed. In meetings, the FOMC makes decisions on a target level for the Federal Funds rate and growth of the U.S. money supply. Each decision includes the future direction of U.S. monetary policy. This study covers all FOMC announcements from 130 scheduled meeting decisions. For the financial crisis period (November 25, 2008 - December 15, 2015), I also include important irregular events related to forward guidance, such as the announcement of LSAP, the chairperson's speech in Jackson Hole and conferences in the dataset.[5]

For each FOMC announcement, I calculate the Fed Funds futures surprise and Treasury futures surprise. Figures 1.4 and 1.5 display the sequence of each surprise. The large fluctuations in the Fed Funds futures surprise in the early 2000s are associated with the Fed's cutting the Fed Funds rate to fight off a recession, terrorist attacks, and the Iraq war. For example, on November 6, 2002, the market expected a 25 basis points cut before the FOMC announcement. However, the Fed decided to lower its Fed Funds rate target by 50 basis points to 1.25 percent. The larger than expected cut led to a big drop in the Fed Funds futures surprise. The next big ups and downs, in 2007 and 2008, correspond to the financial crisis. The sudden drop in Treasury futures on March 18, 2009, implies why I should consider the Treasury futures surprise along with the Fed Funds futures surprise. On this day, there was no change in the Fed Funds rate target. Instead, the Fed announced that it would purchase long-term Treasuries over the next 6 months and increase the size of purchases of agency debt and MBS. The negative value of the Treasury futures surprise reflects the market's response to the Fed's downward pressure on interest rates and forward guidance for the future path of its monetary policy.

[5] I calculate monetary policy surprises for irregular events by using the times for unconventional monetary policy actions provided by Gilchrist et al. (2015b).

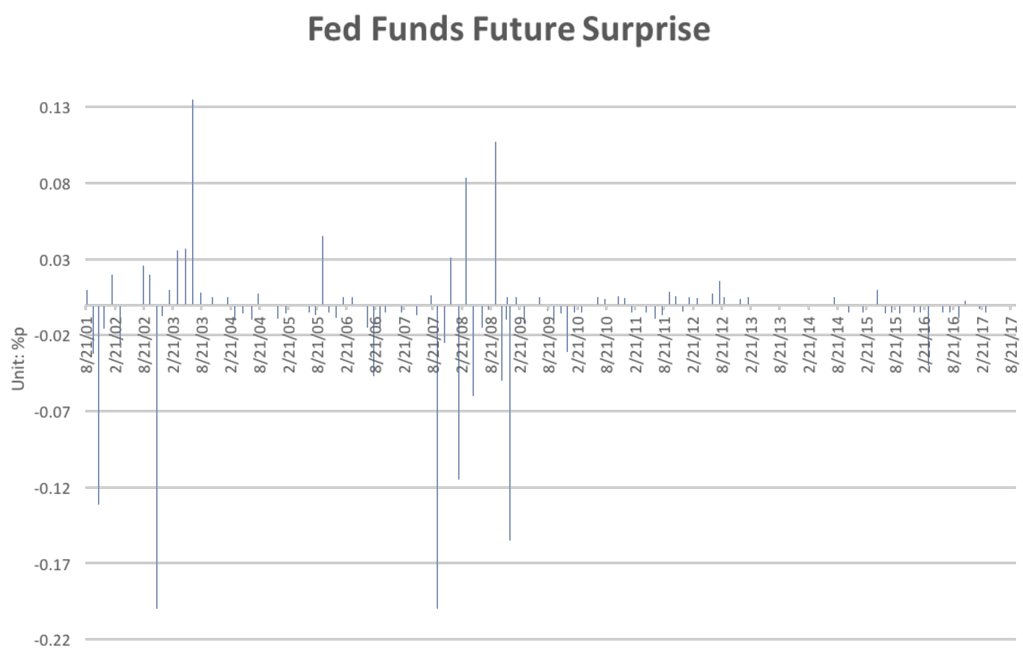


Figure 1.4: Fed Funds Future Surprises (Aug, 2001 - September, 2017)

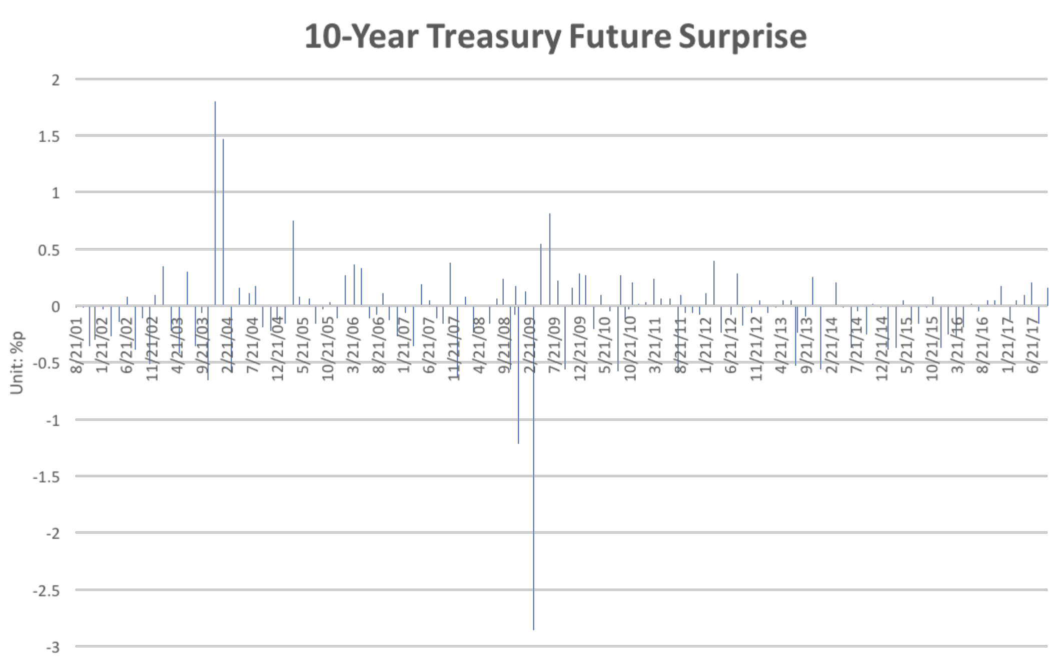


Figure 1.5: Treasury Future Surprises (Aug, 2001 - September, 2017)

1.3.2 Government Bond Yields and Foreign Exchange Rates

For each FOMC meeting and irregular event in the dataset, I collect daily variations in government bond yields and foreign exchange rates for 46 countries. As shown in Table 1.2, the sample countries in my dataset include both developed economies and emerging markets. Changes in an n-year bond yield for country i on FOMC meeting day t within a 1-day period are calculated as

$$\Delta y_{i,t}(n) = y_{i,t}(n) - y_{i,t-1}(n). \tag{1.9}$$

	Division	Country
Europe (23)	Eastern Europe (6)	Bulgaria, Czech Republic, Hungary, Poland, Romania, Russia
	Northern Europe (7)	Denmark, Finland, Ireland, Lithuania, Norway, Sweden, U.K.
	Sothern Europe (4)	Greece, Italy, Portugal, Spain
	Western Europe (6)	Austria, Belgium, France, Germany, Netherland, Switzerland
Asia (13)	East Asia (5)	China, Hong Kong, Japan, South Korea, Taiwan
	South and Southeast Asia (6)	India, Indonesia, Malaysia, Philippines, Singapore, Thailand
	Western Asia (2)	Israel, Turkey
America (7)	North America (2)	Canada, Mexico
	Central and South America (5)	Brazil, Chile, Colombia, Costa Rica, Venezuela
Africa (1)	Africa	South Africa
Oceania (2)	Oceania	Australia, New Zealand

Table 1.2: The Sample Countries

Figure 1.6 depicts the time zone of sample countries. Asian and European markets are closed at the time of the scheduled FOMC announcement. I use the 1-day window between t and $t+1$ for these markets to address a time lag.

The dataset on foreign government bond yield consists of 2-, 5-, and 10-year maturities. I investigate how short-term (2-year), midterm (5-year), and long-term (10-year) yields respond differently to U.S. monetary policy surprises. This allows me to compare the different movements at the short and long ends of the yield curve. To test whether the effects of

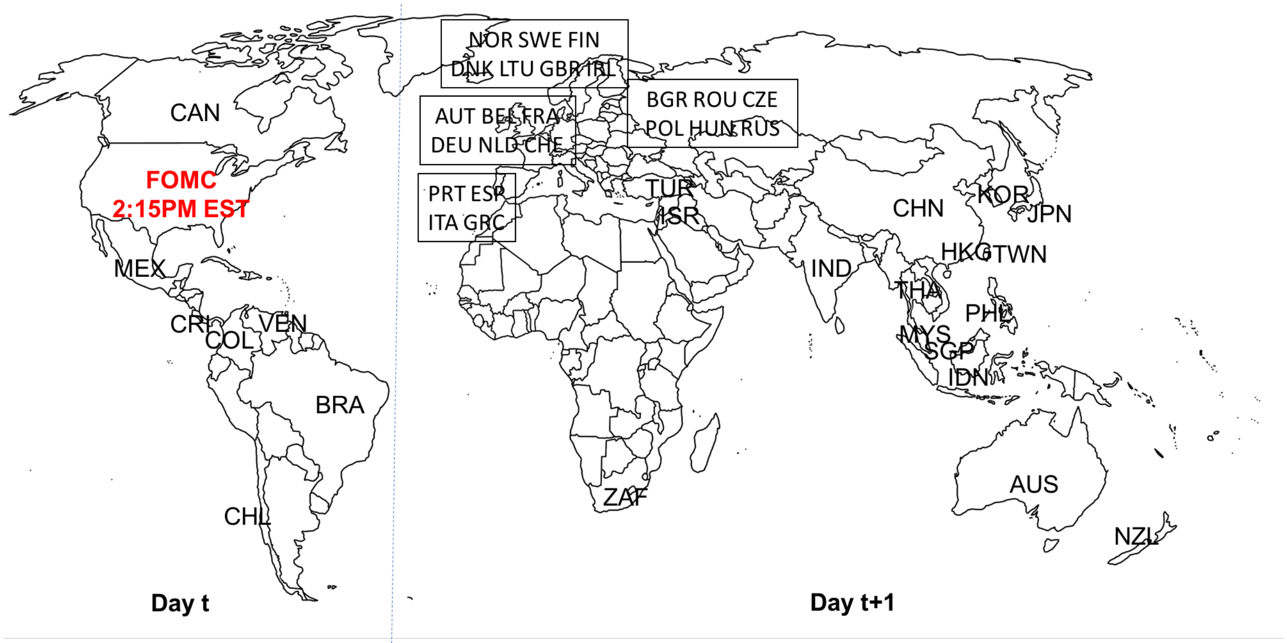


Figure 1.6: Time Zone of Sample Countries

U.S. monetary policy surprises are different across advanced and non-advanced economies, I divide the samples into two groups, developed economies and emerging markets, as shown in Table 1.3.

	Country
Developed Economies	CAD, DEU, FRA, GBR, ITA, JPN, AUT, BEL, NLD, CHE, GRC, PRT, ESP, DNK, FIN, IRL, NOR, SWE, CZE, HUN, POL, KOR, ISR, TUR, MEX, CHL, AUS, NZL
Emerging Markets	LTU, BGN, ROU, RUS, CHN, HKG, TWN, IND, IDN, MYS, PHL, SGP, THA, BRA, COL, CRI, VEN, ZAF

Table 1.3: The Division of Groups

I calculate changes in the foreign exchange spot rate for country i on FOMC meeting day t as follows:

$$\Delta s_{i,t+1} = \frac{s_{i,t+1} - s_{i,t}}{s_{i,t}} \times 100, \tag{1.10}$$

where $\Delta s_{i,t}$ is the percentage changes in the foreign exchange rate (in dollars per unit of non-U.S. currency) within a 1-day window.

The exchange arrangement in each country plays an important role in the responses of exchange rates to U.S. monetary shocks. For example, when a country opens its financial markets to foreign investors, it can experience sudden inflows and stops of foreign funds (Edwards (2007)). A country may fear a floating exchange regime that can magnify their vulnerability to the sudden outflow or inflow of foreign funds. This explains why some countries (mostly emerging markets) are inclined to peg their currency to the U.S. dollar, which may reduce the spillover of U.S. monetary policy surprises. In order to analyze how U.S. monetary policy surprises affect foreign exchange rates under different exchange rate regime, I categorize sample countries into four groups: hard pegs, soft pegs, managed floating, and free floating, as shown in Table 1.4. While most developed economies in my dataset adopt a fully floating exchange regime, many emerging market economies run managed float regimes or limited-flexibility regimes.[6]

	Country
Hard Pegs	LTU, BGR, HKG
Soft Pegs	DNK, CZE, HUN, ROU, RUS, CHN, IND, IDN, MYS, SGP, THA, ISR, CRC, VEF
Managed Floating	CHF, KOR, PHL, TUR, BRA, COL, ZAF
Free Floating	EUR, IRL, NOR, SWE, GBR, POL, JPN, CAN, MEX, CHL, AUS, NZL

Table 1.4: Exchange Rates Arrangement

[6] The exchange rate regime is measured by IMF's *Annual Report on Exchange Arrangement and Exchange Restrictions.*

1.3.3 Empirical Methodology

U.S. monetary policy surprises on FOMC meeting days play a role as exogenous shocks to financial markets in foreign countries. I evaluate the global transmission of U.S. monetary policy surprises to foreign government bond yields and exchange rates using the following panel regression:

$$\begin{aligned}\Delta y_{i,t+1} = \alpha_0 + \beta_1 FF_t + \beta_2 TYF_t + \beta_3 CRISIS + \beta_4 POST + \beta_5 FF_t \cdot CRISIS \\ + \beta_6 TYF_t \cdot CRISIS + \beta_7 FF_t \cdot POST + \beta_8 TYF_t \cdot POST + \mu_i + \varepsilon_{it}.\end{aligned} \tag{1.11}$$

In Equation (1.11), I regress the daily change in country i's financial variables ($\Delta y_{i,t+1}(n)$ for government bond yields and $\Delta s_{i,t+1}(n)$ for exchange rates) around FOMC meeting day t on the Fed Funds futures surprise (FF_t) and Treasury futures surprise (TYF_t). I include $CRISIS$ and $POST$ dummies to identify changes in the influence of U.S. monetary policy surprises after the U.S. financial crisis. $CRISIS$ is 0 in the pre-crisis period (before November 24, 2008) and 1 in the crisis period (i.e., between November 24, 2008, and December 15, 2015). Likewise, $POST$ has the value of 1 in the post-crisis period (after December 15, 2015). I add country fixed effects (μ_i) to capture country-specific time-invariant elements. ε_{it} captures all nonmonetary policy shocks that can affect movement in country i's government bond yields on the FOMC meeting day t.

β_1, β_2, β_3, and β_4 are commonly referred to as the direct effect of FF_t, TYF_t, $CRISIS$, and $POST$ on $\Delta y_{i,t+1}(n)$, respectively. The coefficients β_5, β_6, β_7, and β_8 for interaction terms between monetary policy surprises and dummies help estimate how the effects of monetary policy surprises differ by period.

For example, the net impact of FF_t on $\Delta y_{i,t+1}(n)$ is defined by

$$E[\Delta y_{i,t+1}] = \alpha_0 + \beta_3 CRISIS + \beta_4 POST + (\beta_1 + \beta_5 CRISIS + \beta_7 POST)FF_t. \tag{1.12}$$

The first derivative of Equation (1.12) with respect to FF_t is

$$\frac{\delta E[\Delta y_{i,t+1}]}{\delta FF_t} = \beta_1 + \beta_5 CRISIS + \beta_7 POST. \tag{1.13}$$

In Equation (1.13), β_1 represents the impact of FF_t on $\Delta y_{i,t+1}$ conditional on the value of $CRISIS$ and $POST$ being zero. β_5 indicates whether the effect of FF_t on $\Delta y_{i,t+1}$ is systematically different when $CRISIS$ has the value of 1. For example, a positive β_5 implies that the impact of the Fed Funds futures surprise on the daily change in sovereign bond yields grows more positive in the crisis period compared to the pre-crisis period. Likewise, β_7 allows me to compare differences in the effect of FF_t on $\Delta y_{i,t+1}$ between the pre-crisis and post-crisis period.

Along with the net effect in Equation (1.13), the total effect of FF_t on $\Delta y_{i,t+1}(n)$ in each period is calculated by

$$E[\Delta y_{i,t+1} \mid FF_t \neq 0, CRISIS = 1, POST = 0] = \alpha_0 + \beta_1 + \beta_3 + \beta_5, \tag{1.14}$$

$$E[\Delta y_{i,t+1} \mid FF_t \neq 0, CRISIS = 0, POST = 1] = \alpha_0 + \beta_1 + \beta_4 + \beta_7. \tag{1.15}$$

In Equation (1.14), a positive value of $\alpha_0 + \beta_1 + \beta_3 + \beta_5$ implies that a change in the Fed Funds futures surprise (FF_t) is positively associated with a daily change in foreign government bond yields $(\Delta y_{i,t+1}(n))$ in the crisis period.

1.4 Results

Table 1.5 shows that the response of sovereign bond yields to U.S. monetary policy surprises differs not only across maturities of bonds, but also across periods. For a decrease in the Fed Fund futures surprise of 100 basis points, yields on short-term sovereign bonds in the crisis period would be expected to decline by 80 basis points more than bond yields in the pre-crisis period. A surprise cut in the Fed Fund futures and Treasury futures surprise has a stronger positive association with movement of midterm and long-term sovereign bond yields in the post-crisis period compared to the pre-crisis period. For example, a rise in the Fed Funds futures surprise of 100 basis points leads to an increase of 120 additional basis points in long-term foreign government bond yields in the post-crisis period relative to the pre-crisis period. The Treasury futures surprise also begins to influence the movement of 5-year and 10-year government bond yields in the post-crisis period. For an unanticipated increase in Treasury futures by 100 basis points, foreign government bond yields increase by 5 to 6 additional basis points in the post-crisis period relative to the pre-crisis period. Column (4) in Table 1.5 shows the relationship between foreign exchange spot rates and U.S. monetary policy surprises. My estimates indicate that a decline in the Fed Funds futures surprise of 100 basis points is associated with an appreciation in the local currencies of an additional 9 percent in the crisis period and an extra 16 percent in the post-crisis period, compared to the pre-crisis period.

I attribute these results to the decoupling of interest rates between the U.S. and other countries. In the face of the financial crisis, the Fed lowered the Fed Funds rate to the ZLB. It also cooperated with other central banks to prevent a deepening of the global credit crisis. Although the Fed has continued to raise interest rates since 2015, Europe and Japan have kept their rates near zero, as shown in Figure 1.7. Central banks in emerging markets also did not pursue premature tightening. As a result, the widening gap in interest rates between the U.S. and the rest of the world has caused foreign financial markets to respond sensitively to Fed decisions after the financial crisis.

Table 1.6 shows how sovereign bond yields in a group of developed economies and emerging markets react to U.S. monetary policy surprises. In the crisis period, government bond yields in developed economies significantly respond to unexpected changes in Fed Fund futures across all maturities. For example, the 100 basis points decrease in the Fed Fund futures leads to a drop in government bond yields by 24 to 84 additional basis points in the crisis period, compared to the pre-crisis period. In the post-crisis period, the Treasury

VARIABLES	(1) GOV2	(2) GOV5	(3) GOV10	(4) FX
FF	0.301***	0.307***	0.277***	-1.026***
	(0.0435)	(0.0355)	(0.0411)	(0.235)
TYF	0.0418***	0.0431***	0.0307***	-0.273***
	(0.00974)	(0.0105)	(0.00502)	(0.0379)
FF × CRISIS	0.845***	0.296**	0.400*	-9.089***
	(0.249)	(0.115)	(0.236)	(1.300)
TYF × CRISIS	-0.0278**	-0.00839	0.00669	0.174***
	(0.0123)	(0.0106)	(0.00739)	(0.0620)
FF × POST	-0.198	0.473	1.281**	-16.37***
	(0.282)	(0.285)	(0.496)	(2.689)
TYF × POST	0.0286	0.0534***	0.0656***	-0.116
	(0.0202)	(0.0158)	(0.0162)	(0.124)
CRISIS	0.00199	0.000842	-0.0101***	-0.170***
	(0.00901)	(0.00591)	(0.00364)	(0.0251)
POST	0.00260	0.00271	0.00870*	0.0663***
	(0.0143)	(0.00429)	(0.00458)	(0.0228)
CONSTANT	-0.00291	-0.00275	0.00215	0.0471***
	(0.00362)	(0.00345)	(0.00216)	(0.0118)
Country FE	YES	YES	YES	YES
Observations	4,479	4,436	4,627	4,885
Number of Country	46	45	46	36
Adjusted R-squared	0.00123	0.0293	0.0587	0.0763

NOTE: The dependent variable is daily change in 2-year (GOV2), 5-year (GOV5), 10-year (GOV10) ahead government bond yield and daily percentage change in foreign exchange spot rate in dollars per unit of non US currency (FX) bracketing an FOMC announcement. The entries labeled "FF" denote a 30-minute window change in the Fed Fund Futures around an FOMC announcement. The entries labeled "TYF" denote a 30-minute change in the 10-year Treasury Futures. "CRISIS" is 1 in the sample period between Nov 2008 and Dec 2015. "POST" is 1 in the sample period after Dec 2015. Robust standard errors in parentheses. *** $p<0.01$, ** $p<0.05$, * $p<0.1$

Table 1.5: Response of Government Bond Yields and Exchange Rates to U.S. Monetary Policy Surprises

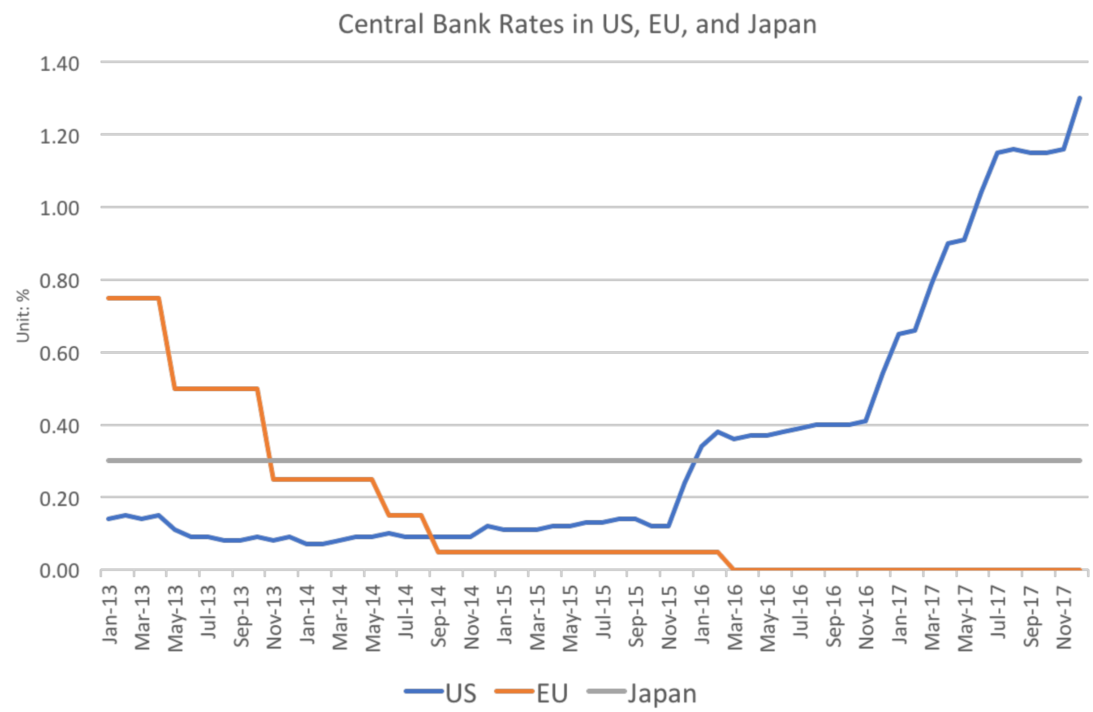


Figure 1.7: Central Bank Rates

futures surprise affects the movement in government bond yields across all maturities. An unexpected increase in Treasury futures by 100 basis points leads to marginal increases in foreign government bond yields by 3 to 7 basis points in the post-crisis period, relative to the pre-crisis period. For emerging market countries, an unanticipated decrease in the Fed Fund futures of 100 basis points is associated with an additional 80 basis points decrease in short-term bond yields in the crisis period, compared to the pre-crisis period. In the post-crisis period, a rise in the Fed Fund futures surprise by 100 basis points is connected to additional increases in midterm and long-term foreign bond yields by 120 to 200 basis points, compared to the pre-crisis period. However, only the response of midterm bond yields shows a statistical significance.

The results suggest that emerging markets' responses to U.S. monetary policy surprises became stronger than those of developed economies after the financial crisis. This finding is consistent with those reported by Chen et al. (2016). Central banks exert greater control

	Developed Economies			Emerging Markets		
	(1)	(2)	(3)	(4)	(5)	(6)
VARIABLES	GOV2	GOV5	GOV10	GOV2	GOV5	GOV10
FF	0.307***	0.325***	0.219***	0.291**	0.265**	0.420***
	(0.0404)	(0.0354)	(0.0287)	(0.123)	(0.0957)	(0.124)
TYF	0.0328***	0.0375***	0.0279***	0.0657**	0.0568*	0.0383***
	(0.00614)	(0.00896)	(0.00588)	(0.0297)	(0.0291)	(0.00934)
FF × CRISIS	0.842**	0.334**	0.247***	0.861*	0.193	0.712
	(0.312)	(0.136)	(0.0799)	(0.433)	(0.214)	(0.739)
TYF × CRISIS	-0.00745	-0.00636	0.0102	-0.0778**	-0.0133	-0.00233
	(0.00995)	(0.0119)	(0.00945)	(0.0344)	(0.0223)	(0.00971)
FF × POST	-0.496**	0.0986	0.809***	0.335	1.218*	2.086
	(0.220)	(0.292)	(0.242)	(0.721)	(0.594)	(1.399)
TYF × POST	0.0329**	0.0726***	0.0738***	0.0207	0.0172	0.0511
	(0.0127)	(0.0163)	(0.0176)	(0.0550)	(0.0310)	(0.0326)
CRISIS	0.00374	-0.00354	-0.0125***	-0.000995	0.0115	-0.00461
	(0.0125)	(0.00725)	(0.00368)	(0.00734)	(0.00965)	(0.00875)
POST	-0.00849	0.00526	0.00692	0.0233	-0.00222	0.0118
	(0.0192)	(0.00632)	(0.00414)	(0.0207)	(0.00350)	(0.0106)
Constant	0.00300	0.00172	0.00480**	-0.0162***	-0.0136*	-0.00386
	(0.00473)	(0.00400)	(0.00220)	(0.00352)	(0.00647)	(0.00542)
Country FE	YES	YES	YES	YES	YES	YES
Observations	3,053	3,029	3,141	1,426	1,407	1,486
Number of Country	28	28	28	18	17	18
Adjusted R-squared	0.000358	0.0265	0.0663	-0.000881	0.0351	0.0551

NOTE: The dependent variable is daily change in 2-year (GOV2), 5-year (GOV5), and 10-year (GOV10) ahead government bond yield bracketing an FOMC announcement. The entries labeled "FF" denote a 30-minute window change in the Fed Fund Futures around an FOMC announcement. The entries labeled "TYF" denote a 30-minute change in the 10-year Treasury Futures. "CRISIS" is 1 in the sample period between Nov 2008 and Dec 2015. "POST" is 1 in the sample period after Dec 2015. Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1

Table 1.6: Comparison of Responses to U.S. Monetary Policy Surprises

over short-term bond yields by their own benchmark interest rates (Caceres et al. (2016)). Monetary policy coordinations on short-term interest rates among developed economies during the financial crisis may explain why the response of 2-year bond yields is greater than those of 5- and 10-year bond yields in the crisis period. On the other hand, long-term bond yields are relatively free to respond to external shocks. For example, the Fed managed to put downward pressure on interest rates under ZLB by purchasing long-term securities. In the post-crisis period, central banks in developed economies are reluctant to raise their short-term target interest rates. This may lead to a larger effect of U.S. monetary policy surprises on the long end of the yield curve rather than the short end. Meanwhile, interest rates around ZLB in developed economies led to cheap borrowing costs in emerging market economies. In the post-crisis period, however, the widening interest rate gap between the U.S. and the rest of the world forced emerging markets to respond sensitively to the tightening U.S. monetary policy.[7]

Table 1.7 shows how the influence of U.S. monetary policy surprises on foreign exchange rates depends on exchange rate arrangements in specific countries. First, there is no exchange rate response to U.S. monetary policy surprises in hard-peg counties. Hard-peg countries, such as Hong Kong, Bulgaria, and Lithuania, have fixed their exchange rates to minimize the vulnerability of their currency to exogenous shocks.[8] In contrast, exchange rates in other regimes significantly respond to unexpected changes in U.S. monetary policy. A surprise decline of 1 percent in Fed Fund futures is associated with an appreciation in local currencies by an additional 6 to 12 percent in the crisis period and an extra 14 to 19 percent in the post-crisis period, compared to the pre-crisis period.

Table 1.8 presents the analysis for how the responses of 10-year government bond yields depend on the exchange rate regime. I find no significant reactions to U.S. monetary policy surprises in hard-pegging countries. However, the movement of interest rates in countries

[7]See Appendix 1 for country-level regressions.

[8]However, a hard-peg country must keep its monetary policy and interest rates in line with the other country. For example, the Hong Kong dollar is pegged to USD, and Bulgaria and Lithuania pegged their currencies to EUR.

Exchange Regime	Hard Peg	Soft Peg	Managed Float	Free Float
VARIABLES	FX	FX	FX	FX
FF	-0.350	-0.768*	-1.043	-1.556***
	(0.152)	(0.377)	(0.787)	(0.338)
TYF	-0.253	-0.133***	-0.407***	-0.379***
	(0.153)	(0.0307)	(0.107)	(0.0674)
FF × CRISIS	-3.363	-6.744***	-12.90***	-11.47***
	(6.056)	(1.744)	(2.934)	(2.446)
TYF × CRISIS	0.103	0.0340	0.396*	0.233
	(0.106)	(0.0656)	(0.186)	(0.131)
FF × POST	-17.97	-14.04***	-19.06*	-18.36***
	(13.07)	(3.364)	(8.462)	(4.988)
TYF × POST	0.491	-0.347**	0.166	-0.121
	(0.316)	(0.139)	(0.451)	(0.198)
CRISIS	-0.160	-0.130***	-0.155*	-0.238***
	(0.0849)	(0.0351)	(0.0682)	(0.0480)
POST	0.0944	0.0146	0.114	0.0910**
	(0.0533)	(0.0356)	(0.0642)	(0.0400)
Constant	0.0363	0.0402**	0.0740*	0.0384
	(0.0299)	(0.0151)	(0.0356)	(0.0244)
Country FE	YES	YES	YES	YES
Observations	408	1,899	951	1,493
Number of Country	3	14	7	11
Adjusted R-squared	0.0461	0.0645	0.0621	0.111

NOTE: The dependent variable "FX" is daily percentage change in foreign exchange spot rate (in dollars per unit of non US currency) bracketing an FOMC announcement. The entries labeled "FF" denote a 30-minute window change in the Fed Fund Futures around an FOMC announcement. The entries labeled "TYF" denote a 30-minute change in the 10-year Treasury futures. Robust standard errors in parentheses. "CRISIS" is 1 in the sample period between Nov 2008 and Dec 2015. "POST" is 1 in the sample period after Dec 2015. Robust standard errors in parentheses.

*** $p<0.01$, ** $p<0.05$, * $p<0.1$

Table 1.7: Response of Foreign Exchange Rate to U.S. Monetary Policy Surprises by Exchange Rate Regime

with a free-floating regime is positively associated with both U.S. monetary policy surprises. For example, a rise in the Fed Fund futures surprise by 100 basis points leads to an increase in the yield by 20 additional basis points in the crisis period and 68 extra basis points in the post-crisis period under the free-floating exchange regime, relative to the pre-crisis period. These results imply that the more flexible exchange arrangement leads to larger magnitudes of responses in sovereign bond yields to U.S. monetary policy surprises. In general, a floating exchange regime magnifies vulnerability to sudden outflows of foreign funds made by carry trade activity in the short run. However, when a country pegs its currency to another or intervenes in exchange markets to stabilize the value of its currency, it can reduce sensitivity to the volatility of capital flow. This explains why hard-pegged exchange regimes do not respond actively to U.S. monetary policy surprises. Also, since my data contain changes within a 1-day window, hard-pegging countries may have a delayed reaction by interest rates to a U.S. monetary policy shock.

1.5 Robustness

1.5.1 Clustering Standard Errors

In my empirical model, government bond yields ($\Delta y_{i,t+1}$) and foreign exchange rates ($\Delta s_{i,t+1}(n)$) change at the country level (i). However, U.S. monetary policy surprises, such as FF_t and TYF_t, vary at the aggregate level, as follows:

$$\begin{aligned}\Delta y_{i,t+1} = \alpha_0 + \beta_1 FF_t + \beta_2 TYF_t + \beta_3 CRISIS + \beta_4 POST + \beta_5 FF_t \cdot CRISIS \\ + \beta_6 TYF_t \cdot CRISIS + \beta_7 FF_t \cdot POST + \beta_8 TYF_t \cdot POST + \mu_i + \varepsilon_{it}. \qquad (1.16)\end{aligned}$$

As a result, I may not assume independence of error terms across countries for each FOMC meeting. The correlation within each FOMC meeting comes from a common error

Exchange Regime	Hard Peg	Soft Peg	Managed Float	Free Float
VARIABLES	GOV10	GOV10	GOV10	GOV10
FF	0.546***	0.328**	0.346	0.230***
	(1.29e-09)	(0.112)	(0.248)	(0.0287)
TYF	0.0770***	0.0460**	0.0282**	0.0255***
	(0)	(0.0161)	(0.0108)	(0.00555)
FF × CRISIS	0.106	0.911	0.206	0.214**
	(0.130)	(0.893)	(0.279)	(0.0898)
TYF × CRISIS	-0.0204	-0.0230	0.0129	0.0193***
	(0.0217)	(0.0284)	(0.00863)	(0.00384)
FF × POST	0.385	1.470	2.767**	0.684***
	(0.895)	(1.585)	(0.919)	(0.166)
TYF × POST	-0.0448	0.0978***	0.0207	0.0727***
	(0.0277)	(0.0247)	(0.0699)	(0.0206)
CRISIS	0.00423	-0.0132	0.0187	-0.0159***
	(0.00212)	(0.0117)	(0.0110)	(0.00292)
POST	0.00154	0.0163	-0.00877	0.00999*
	(0.00218)	(0.0120)	(0.00760)	(0.00495)
Constant	-0.00308***	0.00246	-0.0165*	0.00597***
	(7.68e-05)	(0.00807)	(0.00780)	(0.00149)
Country FE	YES	YES	YES	YES
Observations	207	1,214	606	2,513
Number of Country	3	14	7	21
Adjusted R-squared	0.245	0.0419	0.0525	0.0852

NOTE: The dependent variable is daily change in 10-year (GOV10) ahead government bond yield bracketing an FOMC announcement. The entries labeled "FF" denote a 30-minute window change in the Fed Fund Futures around an FOMC announcement. The entries labeled "TYF" denote a 30-minute change in the 10-year Treasury Futures. "CRISIS" is 1 in the sample period between Nov 2008 and Dec 2015. "POST" is 1 in the sample period after Dec 2015. Robust standard errors in parentheses.
*** p<0.01, ** p<0.05, * p<0.1

Table 1.8: Response of Government Bond Yields to U.S. Monetary Policy Surprises by Exchange Rate Regime

component(ν_t):

$$\varepsilon_{it} = \nu_t + \eta_{it}. \tag{1.17}$$

The nonindependence of error terms (i.e., $\mathbf{E}[\varepsilon_{it}\varepsilon_{jt}] = \rho_\varepsilon \sigma_\varepsilon^2 \neq 0$) may underestimate standard errors with

$$\rho_\varepsilon = \frac{\sigma_\nu^2}{\sigma_\nu^2 + \sigma_\eta^2}, \tag{1.18}$$

which is called a Moulton problem (Moulton (1986)).

I address the possible Moulton problem by clustering standard errors with a block-diagonal in $\hat{\Omega}$:

$$Var(\hat{\beta}) = (X'X)^{-1}X'\hat{\Omega}X(X'X)^{-1}, \tag{1.19}$$

by ordering observations by group.

Table 1.9, with clustering of standard errors, confirms that the global influence of U.S. monetary policy surprises intensified after the financial crisis. A surprise 100 basis point decrease in the Fed Funds futures leads to a drop in government bond yields by 40 to 70 additional basis points across maturities of bonds in the crisis period, compared to the pre-crisis period. In the post-crisis period, for an unexpected rise in Fed Funds futures by 100 basis points, 10-year foreign government bond yields increase by 180 extra basis points, compared to the pre-crisis period. The responses of foreign exchange rates show almost similar results. For an unexpected decrease in the Fed Funds futures by 100 basis points, local currencies appreciate by an additional 9 percent in the crisis period and an extra 16 percent in the post-crisis period, compared to the pre-crisis period.

1.5.2 Isolating the Monetary Policy Surprise Component

In this study, I use two kinds of monetary policy surprises: the Fed Funds futures surprise and the Treasury futures surprise. However, these two surprises may contain overlapping information on the market's response to the Fed's decision, because they are measured within the same time window. I isolate the component of changes in the 10-year Treasury futures

VARIABLES	(1) GOV2	(2) GOV5	(3) GOV10	(4) FX
FF	0.301***	0.307***	0.277**	-1.026**
	(0.104)	(0.112)	(0.112)	(0.491)
TYF	0.0418**	0.0431*	0.0307**	-0.273***
	(0.0196)	(0.0221)	(0.0154)	(0.0788)
FF × CRISIS	0.845***	0.296**	0.400***	-9.089***
	(0.262)	(0.136)	(0.134)	(2.370)
TYF × CRISIS	-0.0278	-0.00839	0.00669	0.174
	(0.0214)	(0.0231)	(0.0172)	(0.185)
FF × POST	-0.198	0.473*	1.281**	-16.37***
	(0.392)	(0.278)	(0.589)	(5.458)
TYF × POST	0.0286	0.0534**	0.0656	-0.116
	(0.0469)	(0.0252)	(0.0448)	(0.717)
CRISIS	0.00199	0.000842	-0.0101	-0.170**
	(0.0119)	(0.00813)	(0.00693)	(0.0670)
POST	0.00260	0.00271	0.00870	0.0663
	(0.0197)	(0.00715)	(0.00967)	(0.101)
CONSTANT	-0.00291	-0.00275	0.00215	0.0471
	(0.00578)	(0.00513)	(0.00407)	(0.0363)
Observations	4,479	4,436	4,627	4,885
Adjusted R-squared	0.00412	0.0337	0.0587	0.0890

NOTE: The dependent variable is daily change in 2-year (GOV2), 5-year (GOV5), 10-year (GOV10) ahead government bond yield and daily percentage change in foreign exchange spot rate in dollars per unit of non US currency (FX) bracketing an FOMC announcement. The entries labeled "FF" denote a 30-minute window change in the Fed Fund Futures around an FOMC announcement. The entries labeled "TYF" denote a 30-minute change in the 10-year Treasury futures. "CRISIS" is 1 in the sample period between Nov 2008 and Dec 2015. "POST" is 1 in the sample period after Dec 2015. Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1

Table 1.9: Clustering and the Responses to U.S. Monetary Policy Surprises

price that is not related to the Fed Funds futures surprise. The isolated component reflects the expected future path of interest rates contained in the FOMC announcement, which is orthogonal to the movement in Fed Funds futures (Wongswan (2009)). I define the isolated surprise component as the *Residual surprise* ($\widehat{Residual}_t$) by the error term from the regression of the Treasury futures surprise on the Fed Funds futures surprise:

$$TYF_t = a_0 + a_1 FF_t + Residual_t. \tag{1.20}$$

Then, I estimate the effects of FF_t and $\widehat{Residual}_t$ on changes in foreign government bond yields $((\Delta y_{i,t+1}))$ and exchange rates $((\Delta s_{i,t+1}))$, as follows:

$$\begin{aligned}\Delta y_{i,t+1} = \alpha_0 &+ \beta_1 FF_t + \beta_2 \widehat{Residual}_t + \beta_3 CRISIS + \beta_4 POST + \beta_5 FF_t \cdot CRISIS \\ &+ \beta_6 \widehat{Residual}_t \cdot CRISIS + \beta_7 FF_t \cdot POST + \beta_8 \widehat{Residual}_t \cdot POST + \mu_i + \varepsilon_{it}.\end{aligned} \tag{1.21}$$

This type of two-step OLS regression with a generated regressor ($\widehat{Residual}_t$) may cause inconsistent estimates of standard errors (Pagan (1984)). To address this problem, I employ a bootstrapping method. Table 1.10 suggests that the responses of government bond yields and exchange rates to Fed Funds futures and Residual surprises become stronger after the financial crisis. For example, an unanticipated decrease by 100 basis points in the Fed Funds futures rate causes foreign government bond yields to decline by 40 to 80 additional basis points in the crisis period, relative to the pre-crisis period. In particular, the Residual surprise plays a significant role in the movement in foreign government bond yields across all maturities after the financial crisis. A hypothetical 100 basis points cut in the Residual surprise leads to an extra 5 to 12 basis points decrease in government bond yields in both the crisis and post-crisis period, compared to the pre-crisis period.

VARIABLES	(1) GOV2	(2) GOV5	(3) GOV10	(4) FX
FF	0.479***	0.498***	0.374***	-1.891***
	(0.0928)	(0.0968)	(0.0447)	(0.343)
Residual	-0.0828**	-0.0930**	-0.0408***	0.372***
	(0.0367)	(0.0406)	(0.0128)	(0.121)
FF × CRISIS	0.783***	0.395***	0.614***	-9.045***
	(0.277)	(0.149)	(0.233)	(1.405)
Residual × CRISIS	0.0967***	0.128***	0.0782***	-0.470***
	(0.0359)	(0.0423)	(0.0135)	(0.127)
FF × POST	0.0398	0.917***	1.827***	-17.34***
	(0.280)	(0.293)	(0.500)	(2.719)
Residual × POST	0.0285	0.0534***	0.0656***	-0.116
	(0.0197)	(0.0151)	(0.0154)	(0.123)
CRISIS	0.00338	0.00101	-0.0108***	-0.171***
	(0.00773)	(0.00644)	(0.00367)	(0.0258)
POST	0.00120	0.000161	0.00558	0.0718***
	(0.0138)	(0.00401)	(0.00441)	(0.0218)
Constant	-0.00487	-0.00451	0.00120	0.0522***
	(0.00475)	(0.00528)	(0.00221)	(0.0180)
Observations	4,479	4,436	4,627	4,885
Number of Country	46	45	46	36
Adjusted R-squared	-0.00886	0.0222	0.0473	0.0655

NOTE: The dependent variable is daily change in 2-year (GOV2), 5-year (GOV5), 10-year (GOV10) ahead government bond yield and daily percentage change in foreign exchange spot rate in dollars per unit of non US currency (FX) bracketing an FOMC announcement. The entries labeled "FF" denote a 30-minute window change in the Fed Fund Futures around an FOMC announcement. The entries labeled "Residual" denote a 30-minute change in the 10-year Treasury Futures that is orthogonal to "FF". "CRISIS" is 1 in the sample period between Nov 2008 and Dec 2015. "POST" is 1 in the sample period after Dec 2015. Robust standard errors in parentheses.
*** p<0.01, ** p<0.05, * p<0.1

Table 1.10: Residual Surprise and the Responses to U.S. Monetary Policy Surprises

1.6 Conclusion

In this study, I investigated how the Fed's dependence on unconventional monetary policy after the financial crisis and its return to conventional policy in 2015 have affected the global influence of U.S. monetary policy. To address this question, I divided sample periods into three phases according to the Fed's monetary policy regimes: pre-crisis, crisis, and post-crisis. I found that the financial crisis significantly strengthened transmission of U.S. monetary policy surprises to foreign government bond yields and exchange rates. My results showed that developed economies became more sensitive to U.S. monetary policy surprises than emerging markets after the crisis.

Overall, my results demonstrate the consequences of the chasm between U.S. monetary policies and those of other countries. While the Fed departed from the ZLB by raising the Fed Funds rate in 2015, central banks in many countries maintained low interest rates and dependence on QE. The global monetary policy divergence forced foreign financial markets to respond elastically to changes in the Fed Funds rate. My findings can help foreign policymakers account for the strengthened influence of post-crisis U.S. monetary policy shocks as they attempt to stabilize their economies.

Chapter 2

U.S. Monetary Policy Shocks, Stock Prices, and Term Spread: Evidence from G7 Countries

2.1 Introduction

Monetary policy plays an important role in the movement of stock and bond markets by affecting the values of private portfolios and the cost of capital (Campbell et al. (2014)). When a central bank raises a nominal interest rate, bond and stock prices go down due to higher borrowing cost. For example, the Dow Jones Industrial Average fell 351.98 points and the broad S&P 500 dropped 1.5 percent when the Federal Reserve (hereafter referred to as the Fed) raised its Fed Funds rate by a quarter point to a target range between 2.25 and 2.5 percent on December 19, 2018. On the same day, the yield on 2-year Treasury bonds rose 4 basis points, while 10-year Treasury notes increased by 2 basis points. As a result, term spreads between short- and long-term bond yields narrowed. In some cases, however, a positive relation between stock and bond prices would break down. For example, the Asian financial crisis in 1997-1998 led to the decoupling of the stock and bond markets in the U.S. (Connolly et al. (2005)). Also, when an increase in the interest rate raises the expected return of the bubble component of assets, stock prices may increase in response to the tightening of monetary policy (Galí and Gambetti (2015)).

In this study, I investigate empirically the influence of monetary policy on stock and bond markets. Specifically, I examine whether the global transmission of U.S. monetary policy surprises to stock price indexes and term spreads in G7 economies changed after the 2007-2008 financial crisis. Using monthly data from 2001 to 2017, I find that a surprising monetary tightening induces a decrease in stock price indexes and term spreads in the pre-crisis period. However, an unexpected monetary tightening also leads to an increase in stock price indexes and term spreads in the post-crisis period.

Policy coordination is important with regard to the global spillover of U.S. monetary policy shocks. For example, the Great Moderation period (mid-1980s to late 2000s) showed that gains from international monetary policy coordination are relatively small compared to the respective monetary policy (Taylor (2013)). However, as Yellen (2012) argues, "persistently strong headwinds restraining recovery and with the Federal Funds rate constrained by the zero bound" under the 2007-2008 financial crisis, international coordination of monetary policy coordination was essential. In the face of the financial crisis, the Fed cooperated with other central banks in developed economies to mitigate the deepening global credit crisis. On October 8, 2008, the Fed and the central banks of the E.U., U.K., and Canada cut their rates by half a point. One week later, the U.S., E.U., and Japan also implemented a coordinated policy to prevent banks from failing. This allows for empirical analysis of how policy coordination during the financial crisis affected the responses of stock and bond markets to U.S. monetary policy shocks.

I calculate U.S. monetary policy shocks by changes in the Fed Funds futures rate between 10 minutes before and 20 minutes after an Federal Open Market Committee (FOMC) announcement. This *monetary policy surprise* measures the unanticipated component of the Fed's decision on the Fed Funds rate target (Kuttner (2001)). Then, for each day of the month, I cumulate surprises on the FOMC date during the previous 31 days. By averaging these monthly surprises across each day of the month (Gertler and Karadi (2015)), I obtain monthly U.S. monetary policy surprises. To analyze the effect of these surprises on stock and bond markets in developed economies, I collect monthly stock price indexes and term spreads between 2-year and 10-year government bond yields in G7 countries. I also consider industrial production and inflation to verify whether estimated results match the expected movement of macro variables based on monetary policy theories.

A vector error correction model (VECM) using monthly data spanning 2001-2017 shows that the influence of U.S. monetary policy on stock price indexes and term premia in G7 countries changes after the financial crisis. Before the crisis, I find that unexpected monetary tightening in the U.S. induces a decrease in stock price indexes and term spreads. This result confirms the conventional view of the relationship between monetary policy and the stock market. A narrowing gap between short- and long-term nominal interest rates also supports the standard view that a tightening monetary policy mainly influences the short end of the yield curve. However, an unanticipated tightening in U.S. monetary policy leads to an increase in stock price indexes and term spreads in the post-crisis period. This positive response of stock prices is at odds with the conventional perception of the effect of monetary policy on the stock market. The existence of a bubble component of asset prices may provide a plausible explanation for the unconventional effect (Galí and Gambetti (2015)). Keeping interest rates close to a zero lower bound for many years in G7 countries led to a lower borrowing cost, which would presumably increase the size of an asset bubble. As a result, the Fed's tapering of quantitative easing and raising the Fed Fund rate since 2015 would lead to a surge in stock prices. A widening of term spreads shows that the impact of an unconventional U.S. monetary policy surprise on the long end of the yield curve is more pronounced in the post-crisis period.

A large literature examines the global spillovers of U.S. monetary policy on foreign stock and bond markets. The literature suggests that monetary policy influences stock market performance through several channels (Bowman et al. (2015); Kurov and Stan (2018); Rey (2016); Tillmann (2016); Wong and Cheung (2016); Bjørnland and Leitemo (2009)). In particular, Wong and Cheung (2016) document that an expansionary U.S. monetary policy leads to higher stock prices in the U.S. and Asia. Likewise, Bjørnland and Leitemo (2009) show that a tightening monetary policy is negatively associated with stock returns. My contribution is to provide empirical evidence that a monetary tightening raises stock prices in a manner consistent with the presence of an asset bubble after the financial crisis.

Turning to the term spread in bond markets, the literature finds that monetary policy affects not only short-term interest rates but also long-term rates (Burger et al. (2017); Gilchrist et al. (2018); Neely (2015); McCauley et al. (2015)). In global financial markets, U.S. monetary policy is an important factor that determines both the short and long ends of the yield curve for dollar lending across borders. For example, Bruno and Shin (2015) show that U.S. monetary expansion leads to more permissive credit conditions in foreign countries. The greater cross-border liability would exert a downward pressure on foreign bond yields. McCauley et al. (2015) find that both emerging markets and developed economies have received large dollar credit inflows since the financial crisis. Bräuning and Ivashina (2019) show that a reduction in the term spread during the financial crisis is associated with more cash flows in dollars to emerging market economies as the Fed depended on quantitative easing under zero lower bound. I show that the term spread widened in advanced economies after the financial crisis.

The remainder of the paper is organized as follows. Section 2.2 discusses the background of the study. Section 2.3 describes the data and methodology used in the paper. Section 2.4 presents results for the spillover estimates of U.S. monetary policy surprises, and Section 2.5 concludes.

2.2 Background

2.2.1 The Financial Crisis and Asset Markets

Table 2.1 outlines the timeline of the financial crisis. In the U.S, the housing boom ended when the subprime mortgage bubble burst in 2007. This was a prelude to the financial crisis. A plethora of foreclosure signs and bailouts and the bankruptcy of Lehman Brothers triggered a panic in global financial markets. In the face of the crisis, the Fed cooperated with other central banks in developed economies to mitigate the deepening global credit crisis. Quantitative easing (QE) by the Fed and the ECB supported lending. Cheap borrowing costs led to the acceleration of asset markets. For example, Figure 2.1 shows that housing prices in the selected G7 countries rose after the financial crisis. Stock price indexes in the U.S., Germany, the U.K., and Japan have also continuously increased since 2009, as shown in Figure 2.2. The uptrend in stock prices overlaps with widening term spreads. This suggests that QE is effective for stimulating the U.S. economy. Given the economic expansion through low interest rates, investors find that asset markets yield higher returns. As a result, demands for safer assets decrease, which may facilitate the growth of the asset bubble.

2.2.2 Rational Asset Price Bubbles

A central bank's decision regarding its key interest rate affects discount rates and, as a result, asset prices. For example, higher interest rates lead to higher bond yields and lower stock prices because the present values of future net cash flows decrease. In addition, a tightening monetary policy may cause investors to expect that a recession would follow, which decreases the market value of firms. Therefore, when a central bank raises short-term interest rates, a frenzy of speculative investment can be restrained.

However, Galí (2014)'s concept of a rational asset price bubble shows that an increase in the interest rate may raise the expected return of the bubble component. According to

Time	Event
Feb 2007	Home sales peak
Mar 2007	Hedge funds housing losses spread subprime misery
Apr 2007	Help for homeowners not enough
Aug 2007	Fed lowers rate to 4.75%
Sep 2007	LIBOR rate unexpectedly diverges
Nov 2007	Treasury creates $75 billion superfund
Dec 2007	Foreclosure rates double
Jan 2008	Fed tries to stop housing bust
Mar 2008	Fed begins bailouts
Apr 2008	Fed lowers rate to 2%
Sep 2008	Lehman Brothers bankruptcy triggers global panic
Sep 2008	Paulson and Bernanke submit bailout to Congress
Oct 2008	Central banks coordinate global action
Nov 2008	Announcement of Large Scale Asset Purchase (LSAP-I)
Dec 2008	Zero interest rates
Nov 2010	Announcement of LSAP-II
Sep 2012	Announcement of LSAP-III
Jun 2013	Announcement of "tapering"
Dec 2015	Fed raises rate for the first time since 2006

Table 2.1: Timeline of the Financial Crisis

this notion, the price in period t of an asset (Q_t) sums to two components: a *fundamental* component (Q_t^F) and a *bubble* component (Q_t^B).

$$Q_t = Q_t^F + Q_t^B \tag{2.1}$$

The fundamental component of an asset price is defined by a discounted stream of payoffs:

$$Q_t^F \equiv E_t\Big\{\sum_{k=1}^{\infty}\Big(\prod_{j=0}^{k-1}\Big(\frac{1}{R_{t+j}}\Big)\Big)D_{t+k}\Big\}, \tag{2.2}$$

where R_t is a riskless real interest rate and D_t is a dividend. The present discounted value of future dividends captures the fundamental component.

On the other hand, the bubble component, defined by the deviation between the asset

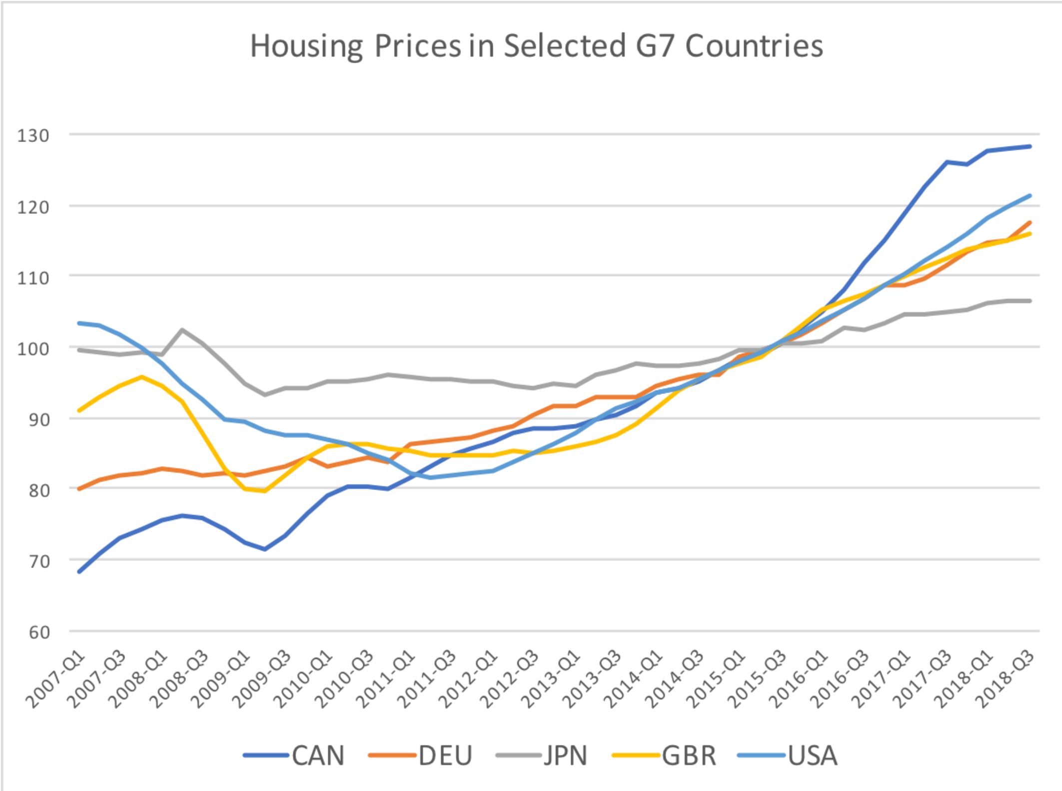


Figure 2.1: Housing Prices in Selected G7 Countries

price and the value of the fundamental component, has no payoffs to discount:

$$Q_t^B R_t = E_t\{Q_{t+1}^B\}. \tag{2.3}$$

When a central bank raises an interest rate, the fundamental value of an asset (Q_t^F) decreases, but the expected growth of the bubble component ($E_t\{Q_{t+1}^B/Q_t\}$) will rise. To formalize the comovement between the bubble innovation and the interest rate surprise, I log-linearize Equation (2.3) and eliminate the expectational operator:

$$q_t^B = q_{t-1}^B + r_{t-1} + \xi_t, \tag{2.4}$$

$$E_{t-1}\{\xi_t\} = 0, \tag{2.5}$$

where lowercase letters are the natural logarithm and $\{\xi_t\}$ is a zero-mean martingale-difference

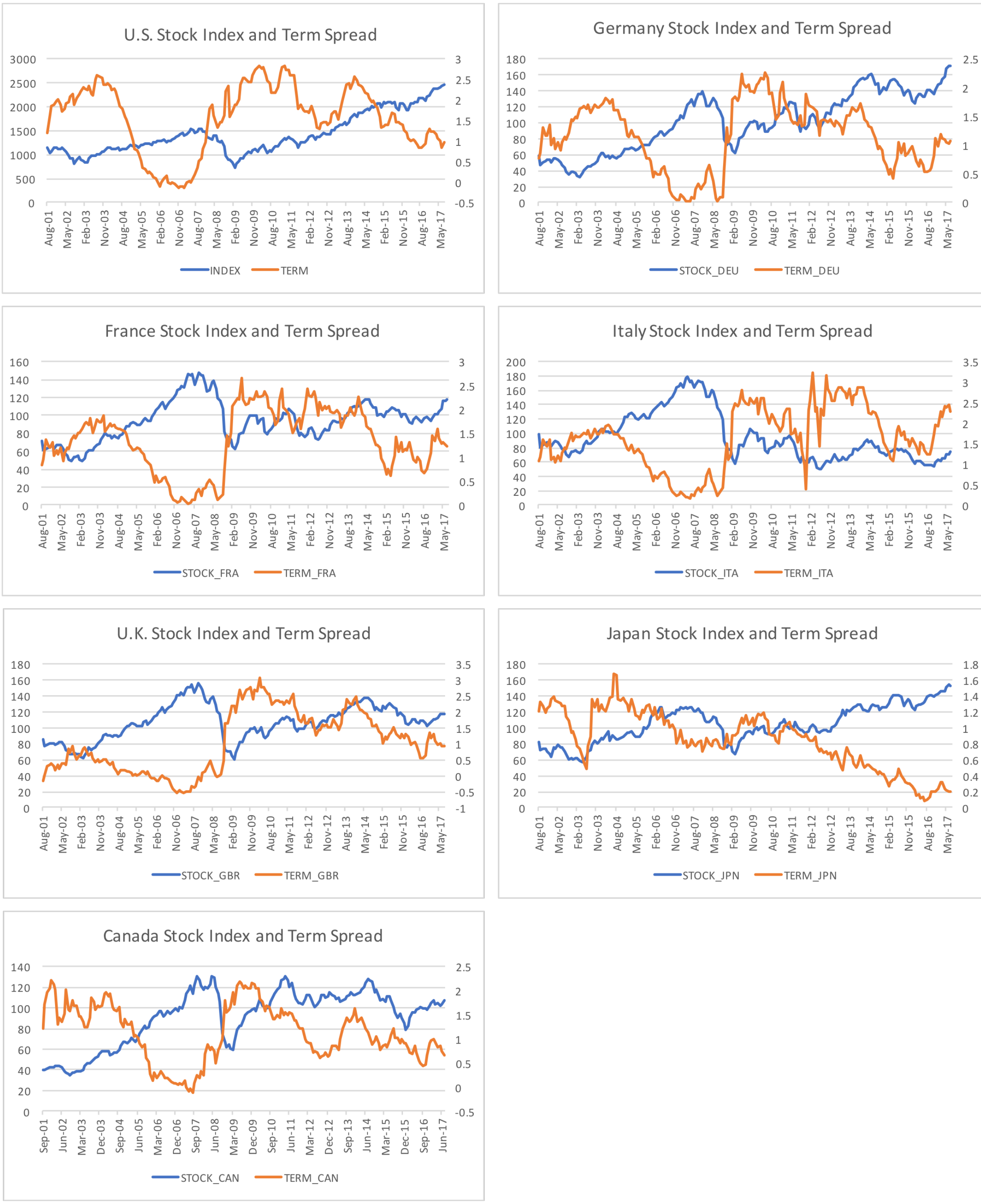


Figure 2.2: Stock Index and Term Spread

process for all t.

I assume that the bubble size is inherently indeterminate, and thus the contemporaneous relationship between an interest rate and the bubble size can be described as follows:

$$\xi_t = \xi_t^* + \psi_r(r_t - E_{t-1}\{r_t\}), \tag{2.6}$$

$$E\{\xi_t^* r_{t-k}\} = 0, \tag{2.7}$$

where $r_t - E_{t-1}\{r_t\}$ is the interest rate innovation and $\{\xi_t^*\}$ is a zero-mean martingale-difference process. $\{\xi_t^*\}$ is orthogonal to all innovations of interest rate, and thus Equation (2.7) holds for $k = 0, \pm 1, \pm 2, \ldots$. As shown in Equation (2.3), a change in interest rate only affects the expected growth rate of the bubble. I assume that the bubble follows an $AR(1)$ process:

$$(1 - \rho_r L)\Delta q_t^B = \varepsilon_{t-1}^r + (1 - \rho_r L)\xi_t^*, \tag{2.8}$$

where $\rho_t \in [0, 1)$, ε_{t-1}^r is an interest innovation, and $\{\xi_t^*\}$ is exogenous relative to the interest rates. Then, an increase in the interest rate has a positive effect on the growth rate of the bubble:

$$\frac{\partial q_{t+k}^B}{\partial \varepsilon_t^r} = \frac{1 - \rho_r^k}{1 - \rho_r} > 0, \tag{2.9}$$

for $k = 0, 1, 2, \ldots$. This leads to a permanent increase in the size of the bubble, given by

$$\lim_{k \to \infty} \frac{\partial q_{t+k}^B}{\partial \varepsilon_t^r} = \frac{1}{1 - \rho_r} > 0. \tag{2.10}$$

According to the model, a tightening monetary policy leads to a larger bubble. This is contrary to the conventional view, that higher short-term nominal interest rates shrink bubbles.

2.2.3 Term Spread

The term spread measures the difference between interest rates at two different maturities. It is most often used to compare of short- and long-term interest rates. The spread is also referred to as the slope of the bond yield curve. A steeper yield curve implies that the gap between short- and long-term interest rates is greater.

The nominal term structure of interest rates can be structured as follows (Morell (2018)):

$$r_{t,(n)} = \frac{1}{n} E_t \sum_{k=0}^{n-1} r_{t+k,(1)}, \tag{2.11}$$

where $r_{t,(n)}$ is an n-period nominal bond yield on time t and $r_{t+k,(1)}$ is the level of the short-term policy rate at time $t+k$. This implies that long-term interest rates are measured by the expected path of future short-term rates.

According to the Taylor rule, the short-term policy rate is determined by several macroeconomic variables such as inflation, output gap, and its first difference (Smets and Wouters (2007)):

$$r_{t,(1)} = \rho_r r_{t-1,(1)} + (1-\rho_r)(r_\pi \pi_t + r_y(y_t - y_t^p)) + r_{\Delta y}(y_t - y_{t-1} - y_t^p + y_{t-1}^p) + \epsilon_t^r \tag{2.12}$$

where π is inflation, y is output, y^p is potential output, and ρ is the degree of interest rate smoothing. r_π, r_y, and $r_{\Delta y}$ imply Taylor rule inflation, output, and Δ output, respectively. Central banks determine the interest rate ($r_{t,(1)}$) in response to the inflation rate and output gap. Combining Equation (2.11) with Equation (2.12) suggests that changes in expected paths for inflation and output in the future affect the term spread. For example, widening spreads precede an economic expansion that makes the yield curve steeper. Narrowing spreads indicate worsening economic conditions, resulting in a flattening of the yield curve. Canada in Figure 2.2 confirms that changes in the term spread can precede the movement of stock prices.

2.3 Method

2.3.1 Data

I use monthly data from August 2001 to September 2017 from G7 countries: Canada, France, Germany, Italy, Japan, the U.K., and the U.S. The variables under consideration are industrial production, consumer price index, term spread (gap between 2-year and 10-year government bond yield), 2-year government bond yield, and stock market indexes: The S&P 500 (U.S.), CAC 40 (France), DAX (Germany), FTSE MIB (Italy), Nikkei 225 (Japan), FTSE 100 (U.K.), and S&P/TSX 60 (Canada). To address the seasonal problem, all variables are adjusted using the X12 approach, as shown in Figure 2.3.

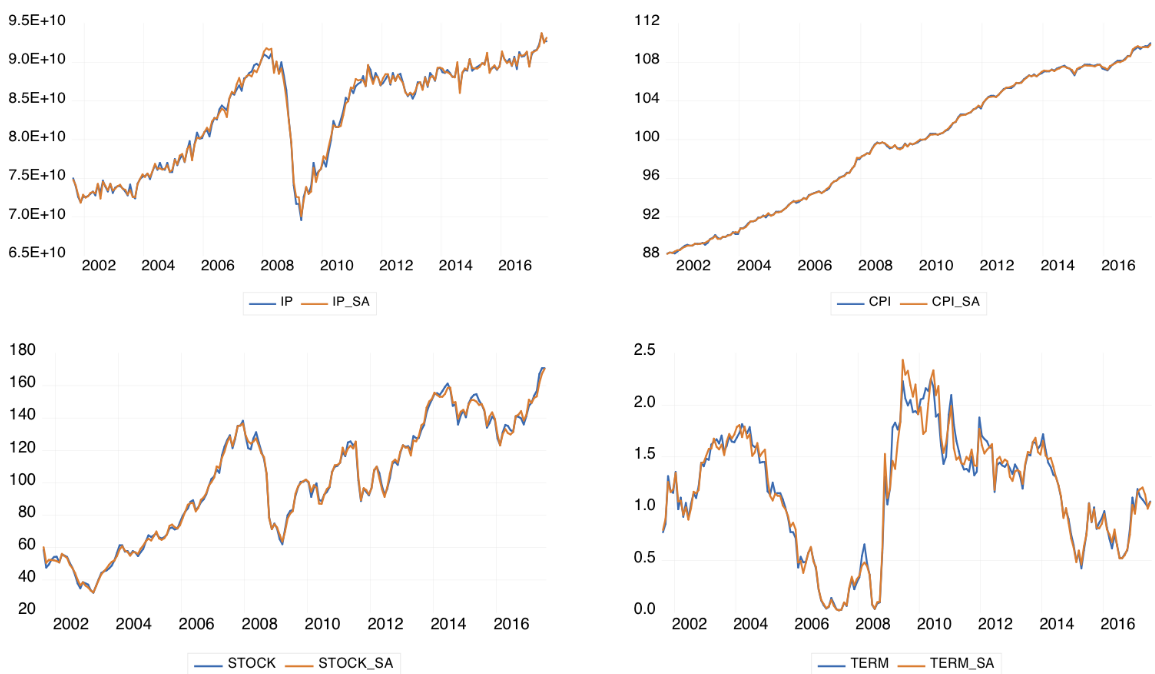


Figure 2.3: Seasonal Analysis of Variables: X12

U.S. monetary policy surprises are calculated by changes in Fed Funds futures rate between 10 minutes before and 20 minutes after an FOMC announcement. This monetary policy surprise measures the unanticipated component of the Fed's decision regarding the Fed Funds rate target (Kuttner (2001)). To extend high-frequency monetary policy surprises to month-level data, I borrow a method from Gertler and Karadi (2015). For each day of the month, I cumulate the surprises on the FOMC date during the previous 31 days. By averaging these monthly surprises across each day of the month, I obtain monthly U.S. monetary policy surprises, as shown in Figure 2.4.

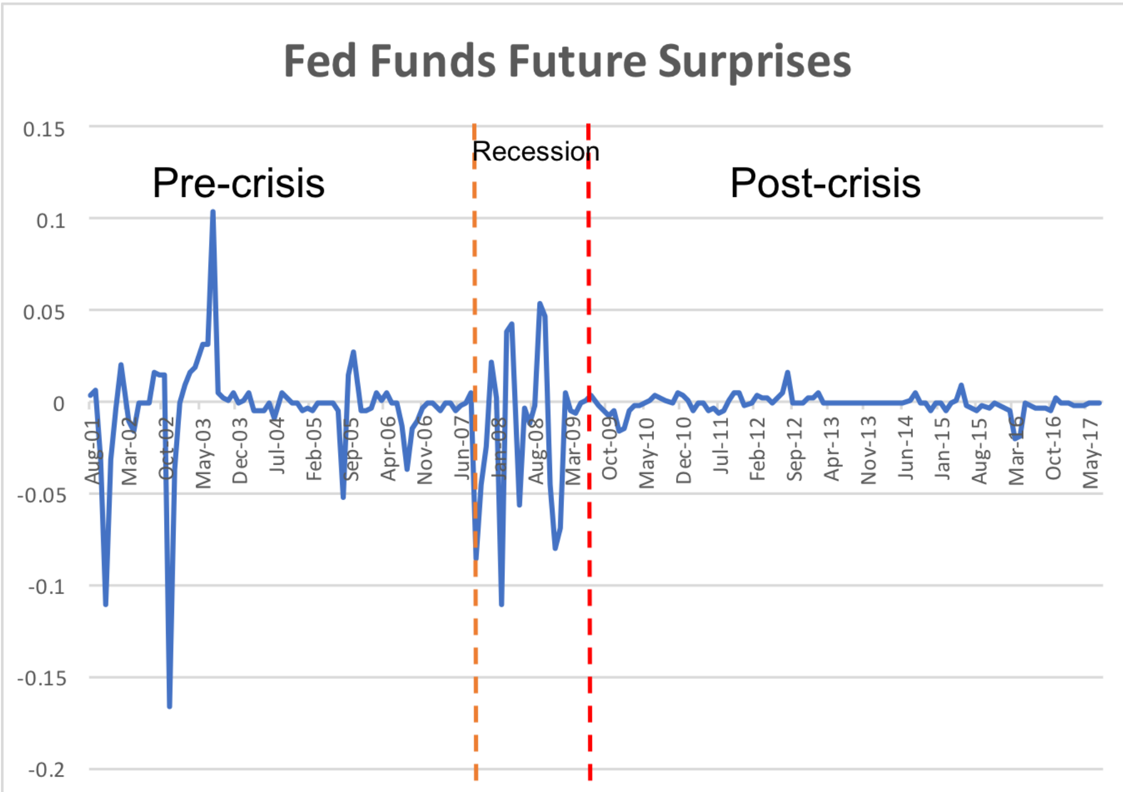


Figure 2.4: Fed Funds Future Surprises

2.3.2 The Vector Error Correction Model

Many macroeconomic variables are nonstationary and drift upward through time. When multiple nonstationary time series share a common trend, they are referred to as cointegrated. A representative vector autoregressive (VAR) is:

$$B\mathbf{y_t} = C(L)\mathbf{y_t} + \varepsilon_t, \tag{2.13}$$

where $\mathbf{y_t}$ is a vector of variables, L is lag operator, and ε_t is a vector of the error term.

To obtain the reduced-form VAR, multiply the Equation (2.13) by B^{-1}:

$$\mathbf{y_t} = A(L)\mathbf{y_t} + e_t, \tag{2.14}$$

where $A(L) = B^{-1}C(L) = A_1L + A_2L^2 + \ldots + A_iL^i$ and $e_t = B^{-1}\varepsilon_t$.

The VECM can be structured when variables in $\mathbf{y_t}$ are nonstationary with unit root ($I(1)$) and linear combinations of the variables are stationary ($I(0)$). I rewrite Equation (2.14):

$$\mathbf{y_t} = (A_1 + A_2 + \ldots)\mathbf{y_{t-1}} - (A_2 + A_3 + \ldots)(\mathbf{y_{t-1}} - \mathbf{y_{t-2}}) - \ldots + e_t. \tag{2.15}$$

By applying the Beveridge-Nelson decomposition, $A(L) = A(1) + (1 - L)A^*(L)$ to Equation (2.15), I obtatin

$$\mathbf{y_t} = (A_1 + A_2 + \ldots)\mathbf{y_{t-1}} - \sum_{i=1}^{\infty} A_i^* \Delta\mathbf{y_{t-i}} + e_t. \tag{2.16}$$

Then, I subtract $\mathbf{y_{t-1}}$ from both sides of Equation (2.16) as follows:

$$\Delta\mathbf{y_t} = A(L)\mathbf{y_{t-1}} - \sum_{i=1}^{\infty} A_i^* \Delta\mathbf{y_{t-i}} + e_t. \tag{2.17}$$

When the rank of $A(L)$ is greater than zero and less than full rank, $A(L)$ can be decomposed as:

$$A(L) = \alpha\beta'. \tag{2.18}$$

The VECM representative form is:

$$\Delta\mathbf{y_t} = \alpha\beta'\mathbf{y_{t-1}} - \sum_{i=1}^{\infty} A_i^* \Delta\mathbf{y_{t-i}} + e_t, \tag{2.19}$$

where the matrix α is a loading matrix that measures how fast errors correct and adjust back to equilibrium. The matrix β is a matrix of cointegration parameters by which the long-run relationship between the variables is governed:

$$\beta' \mathbf{y_t} \sim I(0), \tag{2.20}$$

where the nonstationary time series in the $\mathbf{y_t}$ are cointegrated when a linear combination of the time series is stationary.

In this study, $\mathbf{y_t}$ consists of five variables: U.S. monetary policy surprises (FF), industrial production (IP), consumer price index (CPI), stock index (SP), and term spread ($TERM$).

$$\mathbf{y_t} = [FF, IP, CPI, SP, TERM]'. \tag{2.21}$$

According to Choleski decomposition, the order of the variables in $\mathbf{y_t}$ is expected to have a recursive chain of causality. By putting FF first in the vector in the Equation (2.21), I assume that the Fed makes a monetary decision without considering economic conditions in other G7 countries. However, for the U.S., I place FF after IP and CPI:

$$\mathbf{y_t} = [IP, CPI, FF, SP, TERM]', \tag{2.22}$$

which implies that the Fed accounts for domestic industrial production and inflation when making monetary policy decisions.

I identify whether my time series has a unit root by augmented Dickey-Fuller and Phillip-Perron tests. I find that the random walk of U.S. monetary policy surprises (FF) is stationary; all the other variables are integrated of order one, $I(1)$.[1] Johansen's tests confirm that there are cointegrations among the variables, as shown in Tables 2.2.[2] To choose the number of lags, I use the Akaike Information Criterion (AIC) and the Lagrange Multiplier (LM) test. I also test whether the model has normal distribution, autocorrelation, ARCH, and heteroskedasticity. Based on test results, I conclude that my VECM can be used to examine

[1]Unit root tests are provided in Appendix 2.
[2]Other countries' tests are provided in Appendix 2.

the effect of U.S. monetary policy shocks on stock indexes and term spreads in G7 countries.

Unrestricted Cointegration Rank Test (Trace)

Hypothesized No. of CE(s)	Eigenvalue	Trace Statistic	0.05 Critical Value	Prob.**
None *	0.256300	136.9357	95.75366	0.0000
At most 1 *	0.240273	83.93074	69.81889	0.0025
At most 2	0.097726	34.74215	47.85613	0.4617
At most 3	0.041134	16.33434	29.79707	0.6886
At most 4	0.030591	8.815656	15.49471	0.3827
At most 5	0.018017	3.254423	3.841466	0.0712

Trace test indicates 2 cointegrating eqn(s) at the 0.05 level
* denotes rejection of the hypothesis at the 0.05 level
**MacKinnon-Haug-Michelis (1999) p-values

Unrestricted Cointegration Rank Test (Maximum Eigenvalue)

Hypothesized No. of CE(s)	Eigenvalue	Max-Eigen Statistic	0.05 Critical Value	Prob.**
None *	0.256300	53.00497	40.07757	0.0011
At most 1 *	0.240273	49.18859	33.87687	0.0004
At most 2	0.097726	18.40781	27.58434	0.4615
At most 3	0.041134	7.518685	21.13162	0.9306
At most 4	0.030591	5.561233	14.26460	0.6699
At most 5	0.018017	3.254423	3.841466	0.0712

Max-eigenvalue test indicates 2 cointegrating eqn(s) at the 0.05 level
* denotes rejection of the hypothesis at the 0.05 level
**MacKinnon-Haug-Michelis (1999) p-values

Table 2.2: Johansen's Cointegration Test (Italy)

2.4 Results

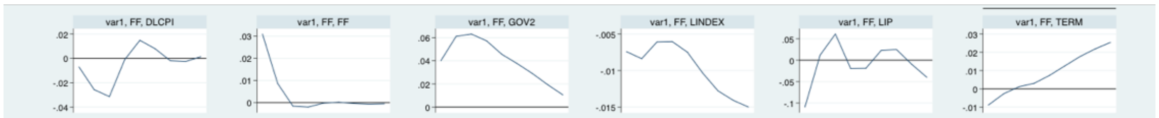
I measure U.S. monetary policy surprises by changes in the response of U.S. financial markets to the Fed's decision. Unexpected changes in U.S. monetary policy play the roles of monetary policy shocks to global financial markets.

I assume that the financial crisis started in December 2007 and ended in June 2009, according to U.S. Business Cycle Expansions and Contractions as defined by the National Bureau of Economic Research (NBER). I divide the sample into two periods: pre-crisis (August 2001 - November 2007) and post-crisis (July 2009 - July 2017). I exclude the financial crisis period (December 2007 - June 2009), because real and financial variables may be too volatile during the crisis, as shown in Figure 2.4.

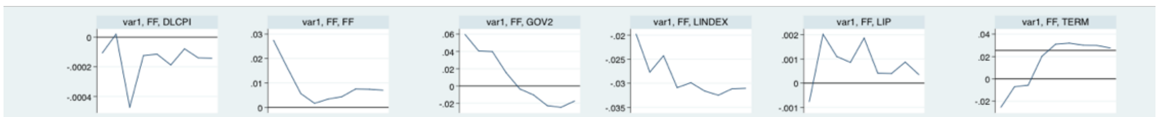
I estimate the effect of an innovation in U.S. monetary policy surprises by 1 standard deviation on industrial production, CPI inflation, 2-year government bond yield, stock index, and term spread. Figure 2.5 presents the impulse response graphs for the effect of positive U.S. monetary policy shocks on real and financial variables before the crisis. Overall results show that a monetary tightening shock leads to an immediate decrease in stock price indexes. This confirms the conventional view of the relationship between monetary policy and stock prices. U.S. monetary policy shocks persistently decrease stock price indexes in France and Italy. In the other countries, however, stock price indexes decrease instantly, then increase over the next few periods. Term spreads also decline following a contractionary monetary policy, except in the U.K. The narrowing spreads due to positive monetary policy surprises imply that worsening economic conditions will follow. Other variables, such as industrial production and CPI inflation, respond to monetary policy shocks as expected: For a positive monetary policy surprise, both industrial production and CPI change negatively.

The impulse response in Figure 2.6 depicts how real and financial variables react to U.S. monetary policy surprises after the crisis. In contrast to the pre-crisis period, a positive monetary policy shock leads to increases in stock price indexes in the post-crisis period. Following an innovation in U.S. monetary policy shocks, the stock indexes in France, Italy, Germany, the U.K., and Japan increase persistently. The positive response of stock prices is at odds with the conventional idea about the effect of monetary policy on the stock market. A plausible explanation for this may be the bubble component of asset prices (Galí and Gambetti (2015)). Keeping interest rates close to a zero lower bound for many years in

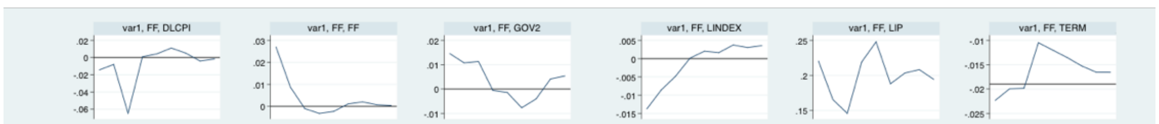
France



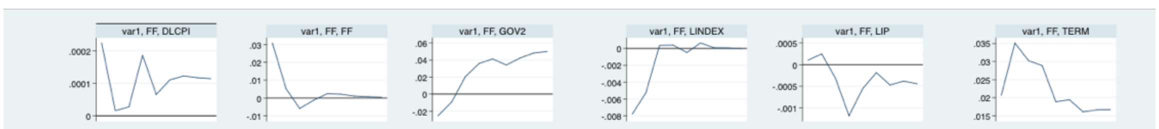
Italy



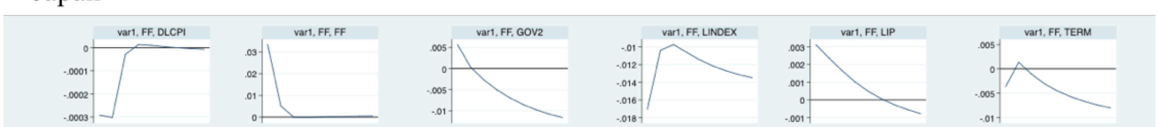
Germany



U.K.



Japan



Canada

U.S.

Figure 2.5: Pre-crisis Impulse Response

G7 countries led to a lower borrowing cost, which would presumably increase the size of an asset bubble. As a result, the Fed's tapering of quantitative easing and raising the Fed Fund rate since 2015 would lead to a surge in stock prices. Monetary tightening is associated with widening term spreads in the post-crisis period. This shows that the impact of U.S. monetary policy surprises on the long end of the yield curve is pronounced in the post-crisis period. An asset bubble would cause investors to obtain higher returns in the stock market. As demands for safer assets decrease, yields on long-term bonds would increase.

2.5 Conclusion

I investigate whether the global transmission of U.S. monetary policy surprises to stock price indexes and term spreads in G7 economies changed after the 2007-2008 financial crisis. Before the crisis, I confirmed the conventional view that U.S. monetary tightening induces a reduction in stock price indexes and term spreads. However, I found that an unanticipated tightening in U.S. monetary policy leads to an increase in stock price indexes and term spreads in the post-crisis period. This demonstrates that zero interest rates for many years in G7 countries led to a lower borrowing cost and increased the demand for asset speculation. As a result, the Fed's tapering of quantitative easing and raising the Fed Fund rates since 2015 made the bubble bigger.

Although I did not cover emerging markets due to data restrictions, I think that cheaper borrowing costs during the financial also led to a bigger asset bubble in emerging market economies. Central bankers in both developed economies and emerging market economies should be cautious about the existence of an asset bubble. Unless they address the size of bubbles by implementing an effective monetary policy, they may experience the sudden collapse of a bubble, which could trigger another financial crisis.

Figure 2.6: Post-crisis Impulse Response

Chapter 3

Exchange Rate and International Trade: The Role of U.S. Monetary Policy

3.1 Introduction

U.S. monetary policy is one of the key factors that determine foreign exchange rates. When the Federal Reserve tightens its monetary policy, U.S. nominal interest rates rise, which in turn induces carry traders more willing to purchase U.S. bonds.[1] The increase in quantity demanded for U.S. bonds induces the appreciation of dollars in the short run.

Foreign exchange rates play an important role in international trade in the sense that a weaker local currency against the dollar stimulates export and improves trade balance.[2] However, a depreciation of a local currency works in a different manner in foreign credit markets. When a country borrows money by dollar-denominated bonds, a stronger dollar may be associated with tighter credit conditions: exporters have to pay higher cost for imported materials, transaction, and shipping. For example, the impact of U.S. monetary tightening propagates to the global value chain. The depreciation of local currencies may lead to higher borrowing costs in dollars, which in turn discourage the production activities of firms in the supply chain. This paper investigates how foreign exchange affects international

[1]See Anzuini and Fornari (2012) for details.

[2]China, India, Hong Kong, and Singapore, for example, which exhibit high trade-to-GDP ratio have intervened the foreign exchange market and sustain their exchange rates low for this reason.

trade through the channel of monetary policy.

To empirically show the effect of real exchange rate on trade through monetary policy, we use monthly data from August 2001 to August 2017 from China, Japan, and Korea. The variables under consideration are the real effective exchange rates, trade volume, trade balance, and unemployment rate. This allows us to compare how large (China and Japan) and small (Korea) open economies in East Asia differently respond to the changes in exchange rates. We assume that U.S. monetary policy is an exogenous shock that determines the real exchange rates for these three countries. We adopt a vector autoregression (VAR) model to demonstrate how the U.S. monetary policy affects trade activity through the real effective exchange rate. We find that the real effective exchange rates decrease in Korea and China, following a tightening U.S. monetary policy. A decrease in real effective exchange rate stands for a gain in trade competitiveness for exporters because exports become more cheap. As a result, an innovation in U.S. monetary policy shocks leads to an immediate increase in the trade volume in Korea. In China, however, real effective exchange rates and trade volumes move to same direction, which is at odd with the conventional view on the relationship between exchange rate and trade.

In order to explain the unconventional responses in China, we introduce a benchmark model in which a representative firm requires multiple production chains and decides the extent of offshoring its task to foreign country (Bruno et al. (2018)). According to the model, a decrease in real effective exchange rates in China limits the total credits which need to be financed to sustain its global value chain (GVC) activity because the production can be discouraged due to higher credit costs incurred by a tightening monetary policy in U.S. In this sences, the opposite responses in Korea and China may be attributed to the different extent of GVC participation between large and small open economies. In general, large open economies participate in the longer supply chain and might be more vulnerable to tightening credit condition.

This paper is in line with the literature which links monetary policy with exchange rate and their impacts on international trade. The relationship between monetary policy and exchange rates has been a subject of interest among the open economy macroeconomics literature (Dornbusch (1976); Rogoff (1996); Eichenbaum and Evans (1995)). Eichenbaum and Evans (1995) investigate the cross-border spillover of U.S. monetary policy. They provide empirical evidences that a contractionary U.S. monetary policy is associated with persistent appreciations in the U.S. real exchange rates. Bruno and Shin (2015) focus on the global spillovers of U.S. monetary policy through the banking sector. They show that a U.S. monetary policy tightening is negatively associated with cross-border bank capital flows, which leads to appreciation of the U.S. dollar. A real depreciation of local currency stimulates exports and improves the trade balance (Arize et al. (2017); Sun and Chiu (2010)).[3] Global value chain activity is an important factor that determines the relationship between exchange rate and trade. Bruno et al. (2018) show that a weaker local currency against the dollar leads to tighter credit conditions and subdue GVC activity, which may lead to a decrease in trade volume. On the other hand, Leigh et al. (2017) do not find any evidences that GVC participation limits the relationship between the exchange rates and trade. The contribution of this paper is that it sheds light on the interaction of monetary policy and exchange rates in determining the trade volume and balance.

The rest of the paper is organized as follows. In Section 3.2, we estimate a VAR model to analyze the effect of U.S. monetary policy on international trade through real exchange rates. in Section 3.3, we interpret the empirical results based on global value chain activity. Section 3.4 concludes.

3.2 Empirical Analysis

We examine empirically how the exchange rate affects international trade. Specifically, we focus on the role of monetary policy. We use monthly data from August 2001 to August

[3]However, lag structure sometimes makes currency depreciation to worsen a trade balance first and to improve it later, which is called as J-curve situation (Bahmani-Oskooee and Ratha (2004)).

	China	Japan	Korea
GDP (trillion, Current USD, 2017)	12.2	4.9	1.5
Total Trade (trillion, Current USD, 2017)	4.6	1.6	1.2
Trade to GDP ratio (%)	37.8	31.2	80.8
Trade with U.S. (biilion, Current USD, 2016)	710.4	283.6	144.6
Trade with U.S. to GDP ratio (%)	6.3	5.7	10.2
Trade with U.S. to Total Trade Ratio (%)	15.3	17.3	11.6
De Jure Financial Openness	-1.19	2.39	0.34

Source: GDP and Trade data from World Bank, De Jure Financial Openness from the Chinn-Ito Index

Table 3.1: Sample Country Characteristics

2017 from China, Japan, and Korea. This allows us to compare how differently large open economies (China and Japan) and small open economy (Korea) in East Asia respond to the changes in exchange rates. As shown in Table 3.1, the sample countries show their own characteristics regarding the trade. For example, Korea's high trade to GDP ratio implies that its economy intensively depends on trade activity and is vulnerable to exogenous economic shocks. Comparing to Korea, trade to GDP ratio is relatively low in China and Japan. However, the portions of trade with U.S. among their total trade are bigger in China and Japan than in Korea.

The sample countries' trade integrations with U.S. allows us to analyze how U.S. monetary policy influences trade volume and trade balance through the channel of real exchange rate. We assume U.S. monetary policy is an exogenous variable that determines the real exchange rates in these East Asian countries.

U.S. monetary policy surprise is calculated by changes in the Fed Funds futures rate between 10 minutes before and 20 minutes after an FOMC announcement. This monetary policy surprise measures the unanticipated component of the Fed's decision on the Fed Funds rate target (Kuttner, 2001). To extend the high frequency monetary policy surprises to the monthly level data, we borrow a method from Gertler and Karadi (2015). For each day of the month, we cumulate the surprises on FOMC date during the last 31 days. By averaging these monthly surprises across each day of the month, we obtain monthly U.S. monetary policy surprises as Figure 3.1.

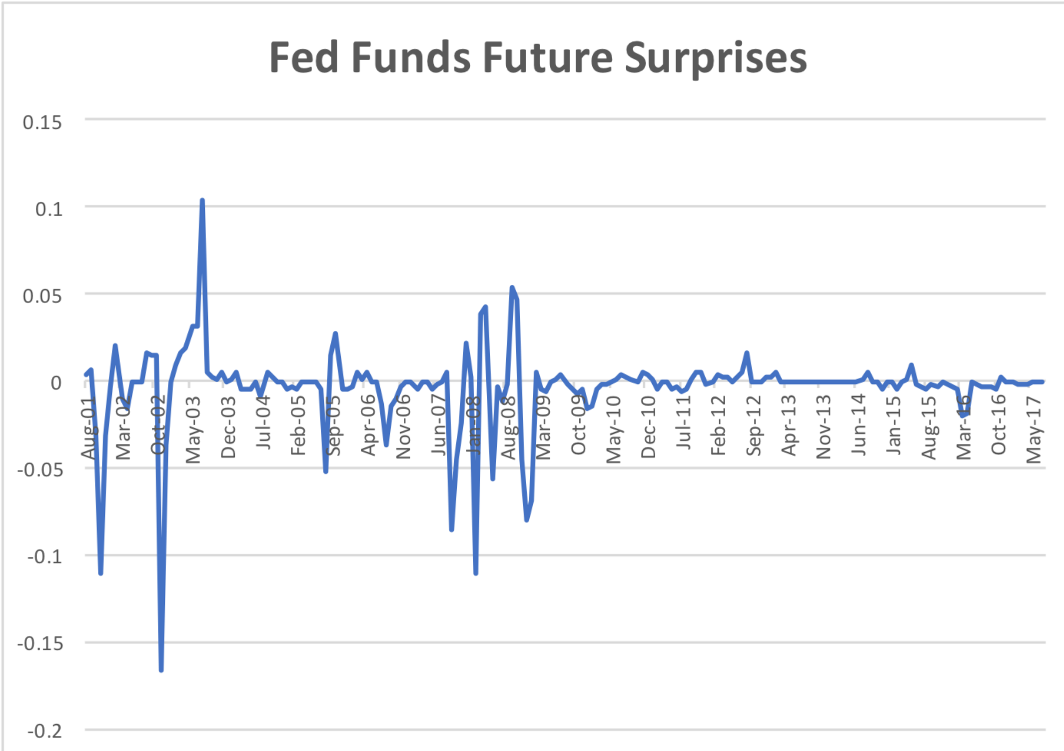


Figure 3.1: U.S. Monetary Policy Surprises

3.2.1 Empirical Framework

The Vector Autoregression (VAR) allows us to analyze the dynamic response of the variables of interest to external shocks. A representative VAR can be expressed as

$$By_t = C(L)y_t + \varepsilon_t, \tag{3.1}$$

where y_t is a vector of endogenous variables, B and C are matrices of the estimated coefficients, L is a lag operator, and ε_t is an error term that is I.I.D. The reduced form of Equation (3.1) is

$$y_t = A(L)y_t + v_t, \tag{3.2}$$

where $A(L) = B^{-1}C(L) = A_1L + A_2L^2 + \cdots + A_iL^i$, $v_t = B^{-1}\varepsilon_t$, and i is the number of lag.

Equation (3.2) can be expressed as

$$y_t = \frac{1}{[I - A(L)]} v_t = K(L) v_t, \tag{3.3}$$

which allows for the estimation of the impulse response and variance decomposition functions. We assume that the B^{-1} is a lower-triangular matrix and, thus, we can identify the residuals with the Choleski decomposition as follows:

$$E(v_t v_t^{'}) = (B^{-1}\varepsilon_t)(B^{-1}\varepsilon_t)^{'} = (B^{-1})\varepsilon_t \varepsilon_t^{'}(B^{-1})^{'} = (B^{-1})(B^{-1})^{'}, \;\; E[\varepsilon_t \varepsilon_t^{'}] = I_n. \tag{3.4}$$

In the recursive form of the model with a lower-triangular B^{-1}, the order of the variables is important. For example, the first variable in the system affects the innovation of other variables below it. However, the other variables cannot affect contemporaneously the innovations of the variables above them in the sequence. Therefore, the order of variables should be determined based on the recursive chain of causality. We develop a VAR system with the ordering: U.S. monetary policy surprises, real effective exchange rate, trade volume (trade balance), and unemployment rate.

We adjust variables by the $X12$ approach to address the seasonal problem. For example, the real effective exchange rate in China has a seasonal characteristic as shown in Figure 3.2. In this case, we use the seasonally adjusted values instead of original values.

In order to test if the variables are stationary, we exploit Augmented Dickey-Fuller and Phillips-Perron tests. For example, the real effective exchange rates, trade volume, and import are found to be integrated with $I(1)$ the ADF unit roots test in Table 3.2. Phillips-Perron tests also confirm this as shown in Table 3.3.

We also conduct the Johansen's cointegration test. Table 3.4 implies that variables in China have long-run relationships by having $I(1)$. Thus, we employ the VEC model to address the cointegraing vectors for Chinese dataset. To choose the number of lags, we use the Akaike Information Criterion (AIC) and the Lagrange Multiplier (LM) test. And we also test

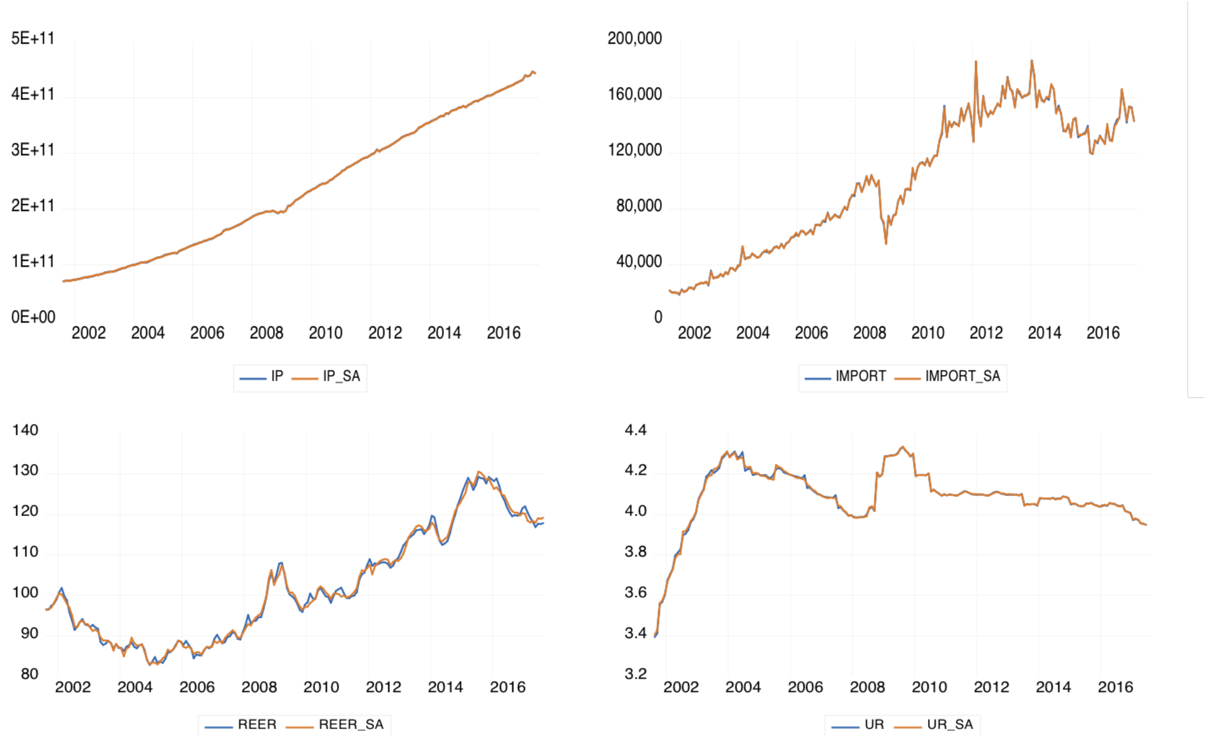


Figure 3.2: Seasonal Analysis of Variables (China)

Variables	Level	First difference	Integration order
Real Effective Exchange Rate	-0.768365	-9.054669	I(1)
Trade Volume	-1.604605	-13.34692	I(1)
Import	-1.321702	-15.59871	I(1)
Unemployment Rate	-3.301649	-4.804930	I(0)

Table 3.2: Augmented Dickey–Fuller Tests (China)

Variables	Level	First difference	Integration order
Real Effective Exchange Rate	-0.609547	-9.076104	I(1)
Trade Volume	-2.060844	-21.52221	I(1)
Import	-1.381559	-21.58880	I(1)
Unemployment Rate	-4.833360	-14.37547	I(0)

Table 3.3: Phillips-Perron Tests (China)

Unrestricted Cointegration Rank Test (Trace)				
Hypothesized No. of CE(s)	Eigenvalue	Trace Statistic	0.05 Critical Value	Prob.**
None*	0.194872	77.78834	47.85613	0.0000
At most 1*	0.098151	37.47203	29.79707	0.0054
At most 2*	0.078617	18.25681	15.49471	0.0187
At most 3	0.016144	3.027314	3.841466	0.0819

Unrestricted Cointegration Rank Test (Maximum Eigenvalue)				
Hypothesized No. of CE(s)	Eigenvalue	Max-Eigen Statistic	0.05 Critical Value	Prob.**
None*	0.194872	40.31630	27.58434	0.0007
At most 1	0.098151	19.21522	21.13162	0.0908
At most 2*	0.078617	15.22950	14.26460	0.0351
At most 3	0.016144	3.027314	3.841466	0.0819

Notes: Trace test indicates 3 cointegrating equations at the 0.05 level. The Max-egenvalue test also indicates 1 conintegrating equation at the 0.05 level. * denotes regions of the hypothesis at the 0.05 level. ** illustrates MacKinnon-Haug-Michelis (1999) p-values.

Table 3.4: Johansen's Cointegration Test (China)

Unrestricted Cointegration Rank Test (Trace)				
Hypothesized No. of CE(s)	Eigenvalue	Trace Statistic	0.05 Critical Value	Prob.**
None*	0.199742	61.64961	47.85613	0.0015
At most 1	0.057679	19.98213	29.79707	0.4241
At most 2	0.034727	8.872673	15.49471	0.3773
At most 3	0.012030	2.263303	3.841466	0.1325

Unrestricted Cointegration Rank Test (Maximum Eigenvalue)				
Hypothesized No. of CE(s)	Eigenvalue	Max-Eigen Statistic	0.05 Critical Value	Prob.**
None*	0.199742	41.66748	27.58434	0.0004
At most 1	0.057679	11.10946	21.13162	0.6364
At most 2	0.034727	6.609370	14.26460	0.5363
At most 3	0.01230	2.263303	3.841466	0.1325

Notes: Trace test indicates 1 cointegrating equations at the 0.05 level. The Max-egenvalue test also indicates 1 conintegrating equation at the 0.05 level. * denotes regions of the hypothesis at the 0.05 level. ** illustrates MacKinnon-Haug-Michelis (1999) p-values.

Table 3.5: Johansen's Cointegration Test (Korea)

whether the model has normal distribution, autocorrelation, ARCH, and heteroscedasticity.

3.2.2 Results

Figure 3.3 shows the impulse responses of the variables (real effective exchange rate, trade volume, and unemployment rate) in Korea. Following positive U.S. monetary policy surprises, the real effective exchange rates decrease. This indicates that monetary tightening leads to higher money inflow to U.S. and the appreciated value of dollars. As a result, the currencies in other countries are more likely to depreciate relatively. A decreasing real effective exchange rate stands for a gain in trade competitiveness because exports become more cheap. This explains what happens to trade volume. An innovation in U.S. monetary policy shocks leads to an immediate increase in the trade volume in Korea. However, it decreases after 2 periods.

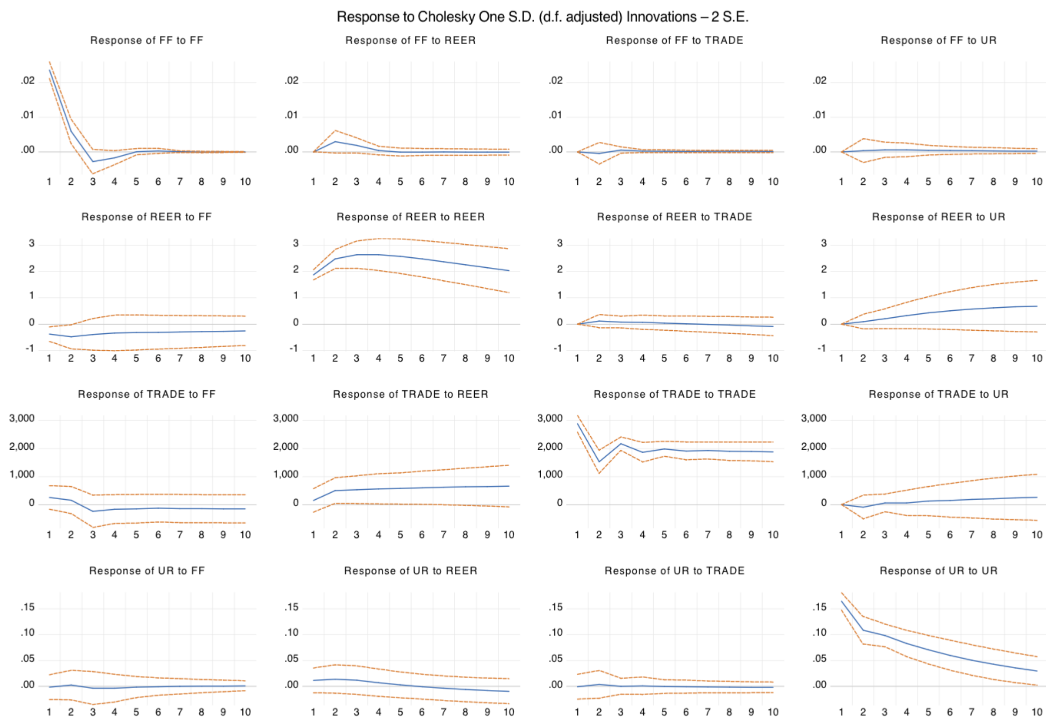


Figure 3.3: Korea's Impulse Response to U.S. Monetary Policy Surprises (VAR) - Trade Volume

In Figure 3.4, the impulse responses in China show similar results. Real effective exchange rates increase slight initially but they dropped after the second period. Likewise, the initial positive response of trade volumes to U.S. monetary policy shock changes negatively later. This is at odd with the conventional view on the relationship between exchange rate and trade. In general, a higher real effective exchange rate indicates a loss in trade competitiveness and, thus, lower trade volume.

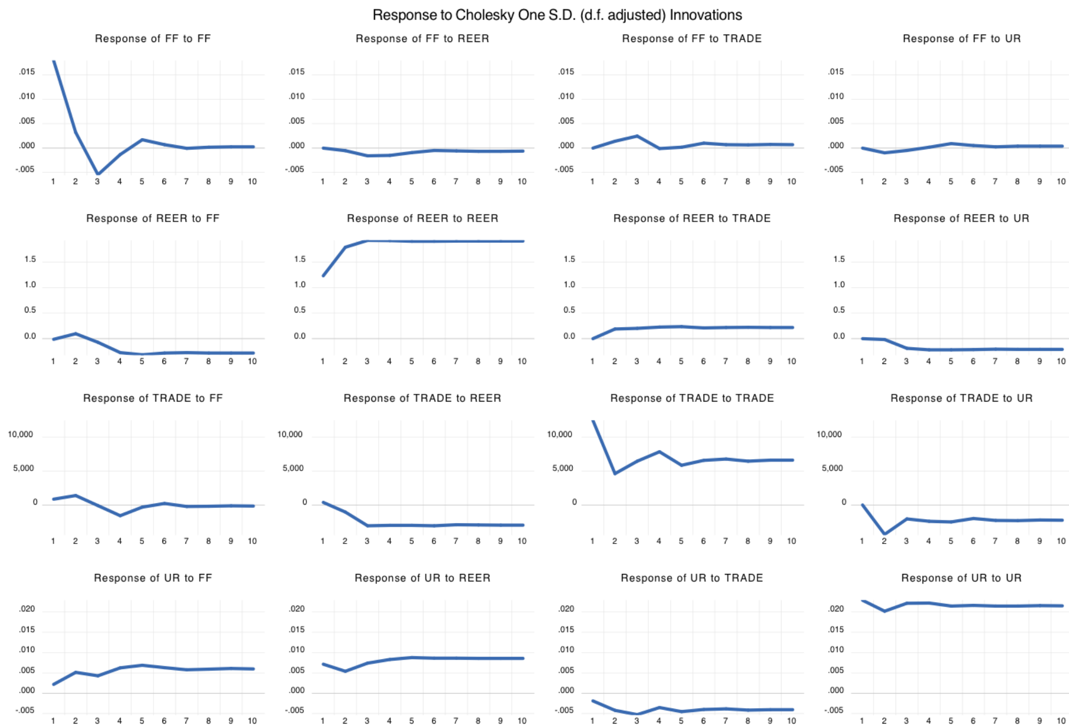


Figure 3.4: China's Impulse Response to U.S. Monetary Policy Surprises (VECM) - Trade Volume

Figure 3.5 depicts how the variables respond to U.S. monetary policy shocks in Japan. Overall, the level of responses in Japan are smaller than other two countries. There are no instant reactions in the real effective exchange rates and trade volumes. Then, the trade volume increases as the local currency appreciates. Although the real exchange rates decreases later, the trade volume shows a positive response to a contractionary U.S. monetary policy.

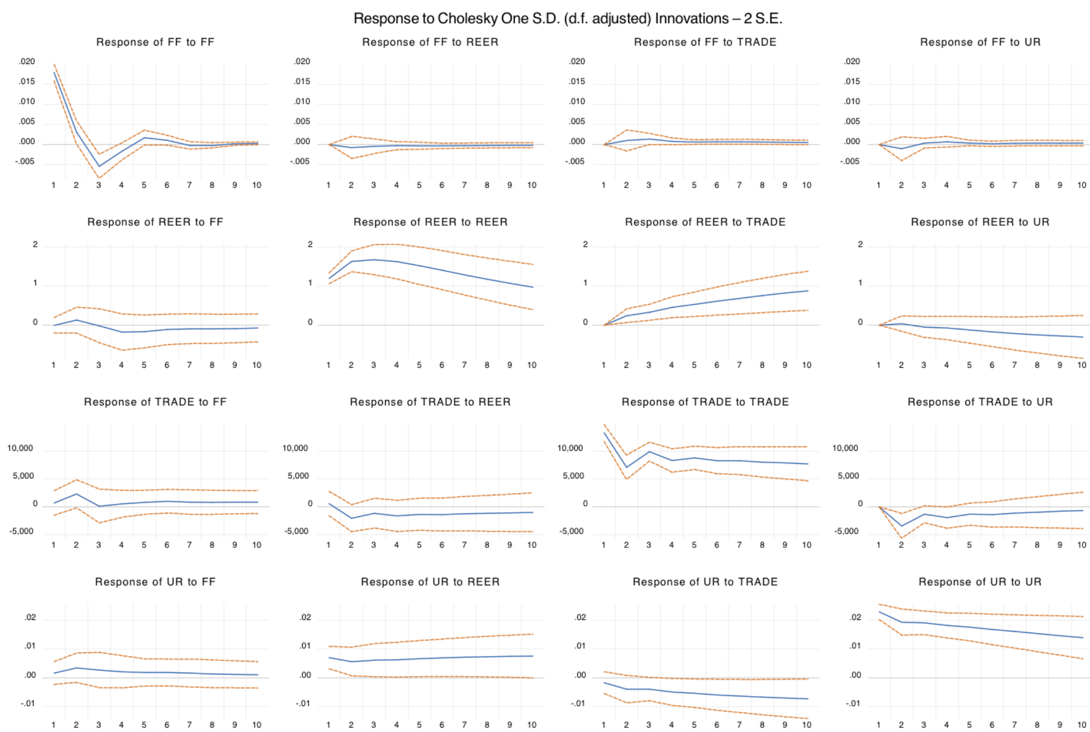


Figure 3.5: Japan's Impulse Response to U.S. Monetary Policy Surprises (VAR) - Trade Volume

As shown in Figure 3.6, Korea's response to U.S. monetary policy surprises confirms a conventional view on the relationship between real exchange rate and trade balance. A contractionary monetary policy in U.S. is associated with a depreciation of local currencies in Korea. The devaluation of currency leads to an improvement of the trade balance in Korea.

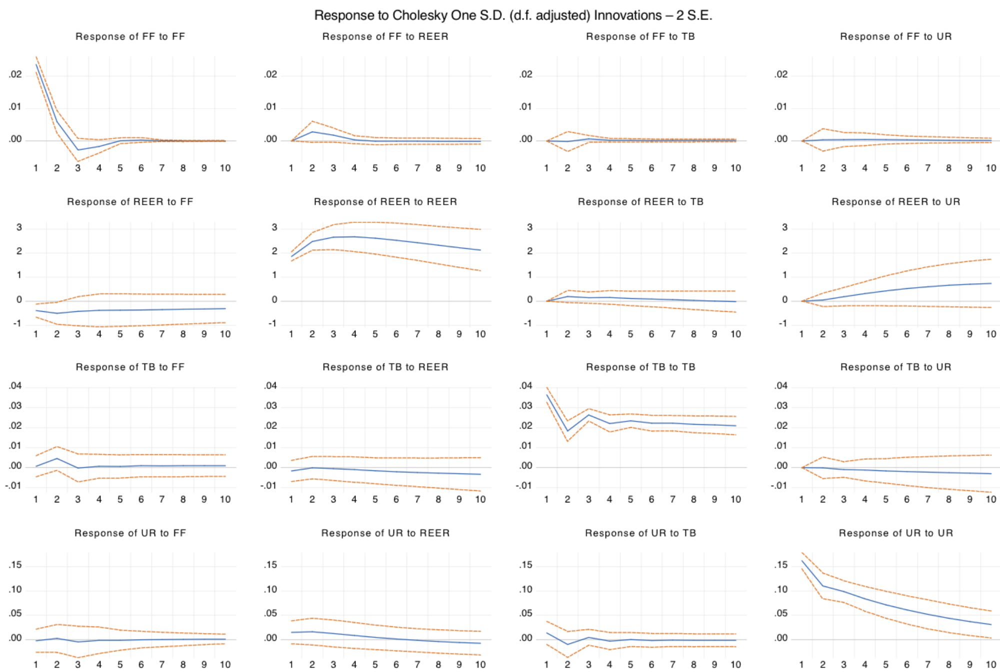


Figure 3.6: Korea's Impulse Response to U.S. Monetary Policy Surprises (VAR) - Trade Balance

However, Figure 3.7 shows that China responds to U.S. monetary policy differently. The real exchange rate in China reacts positively to a monetary tightening in U.S. An increase in real exchange rate implies a loss in trade competitiveness because exports become more expensive. As a result, the appreciation of local currency worsens the trade balance in China.

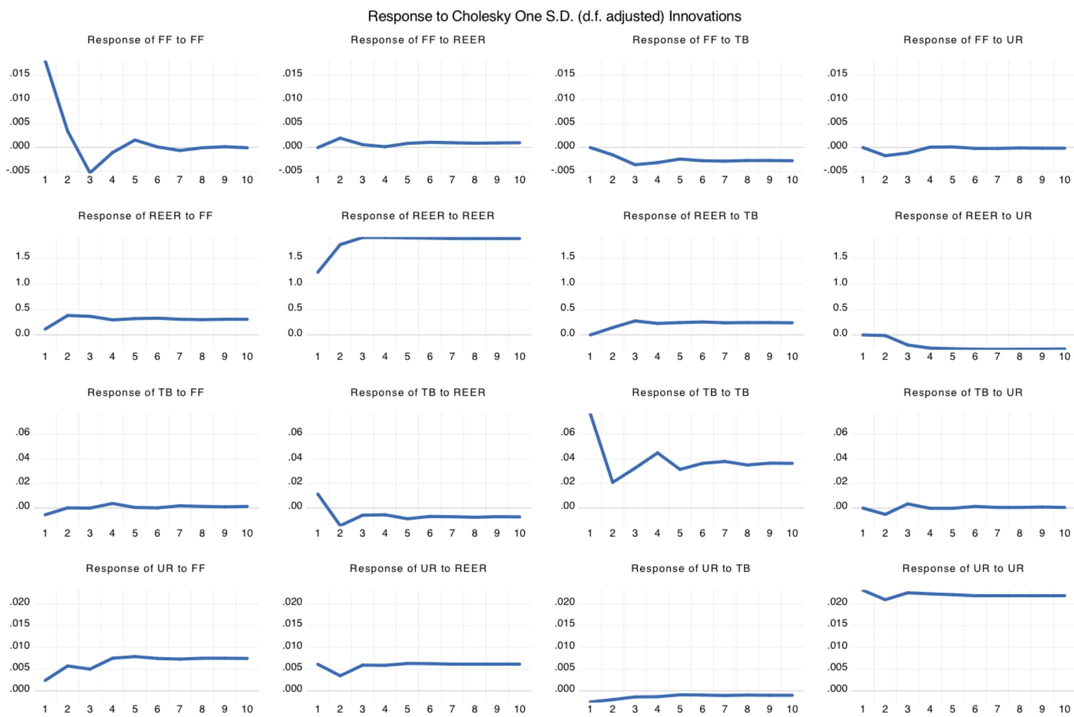


Figure 3.7: China's Impulse Response to U.S. Monetary Policy Surprises (VECM) - Trade Balance

3.3 Discussion

3.3.1 Global Value Chain

In the previous empirical analysis, we find that Korea and China respond differently to U.S. monetary policy surprises. U.S. monetary tightening leads to a decrease in real effective exchange rates in both countries. As a result, Korea's trade volume raises as exports become more cheap. In China, however, a decrease in the real effective exchange rate is followed by lower trade volumes. Likewsie, a contractionary U.S. monetary policy improves the trade balance in Kora but deteriorate it in China.

The GVC activity may provide a plausible explanation for China's unconventional response to U.S. monetary policy. Bruno et al. (2018)'s offshoring models show that the higher credit costs due to the depreciation of local currencies would lead to lower level of offshoring in the global value chain. From the fact that trade for intermediate goods accounts for two-thirds of total trade (Bems et al., 2011), the changes in exchange rate and offshoring would influence total trade volume and trade balance.

For example, a decrease in the real effective exchange rate may increase a borrowing cost in dollars. The depreciation of local currency limits the total credits which need to be financed to sustain the global value chain since the production activities can be discouraged due to higher credit costs incurred by U.S. monetary policy. The opposite responses in Korea and China may be attributed to the different extent of GVC participation between large and small open economies. In general, large open economies participate in the longer supply chain and might be more vulnerable to tightening credit condition.

3.3.2 Production Chains with Offshoring and Emissions

The role of global value chain as the linkage of monetary policy and international trade allows us to extend the benchmark model.[4] Our extended model describes the offshoring

[4]The extended model is provided in Appendix 3.

behavior of the representative firm in response to environmental regulations and higher credit costs arising from a monetary policy favoring dollars. We show that the greater credit costs are, the less likely the representative firm is willing to offshore particularly when the environmental regulations are loose.

We also demonstrate the possibility that countries with sufficiently stringent environmental standards may keep offshoring production stages abroad even when they face high credit costs in order to minimize emission tax penalties. These findings imply that monetary policies may possess differential impacts on the environmental quality depends upon the stringency of environmental regulations across countries in the production supply chain. Due to the relative significance of environmental policies over tightening credit conditions in the North countries, stronger dollars may harm less or even improve the air quality in the North. However, the effect of the monetary policies on the environmental quality are likely to outweigh that arising from environmental regulations in the South. Stronger dollars strictly induce the South to experience lowered air quality.

Figure 3.8 shows that the CO2 emissions have been decreased in the G7 countries (i.e., the representative North Countries). One of the plausible reasons for how they can sustain economic growth while keeping stringent environmental standards is offshoring emission generating productions, as suggested by our model. On the other hand, Figure 3.9 illustrates that CO2 emissions have been increased in the selected emerging market economies (i.e., the representative South Countires). Relatively lenient environmental standards might restrict the offshoring the emission intensive productions and leads to higher emissions in these countries. At this point, however, it is limited to identify the role of offshoring in GVC in the total emission levels.

Our future work will focus on showing how a change in credit conditions due to monetary policy determines the extent of offshoring emission generating productions in global value chain. To confirm the arguments of our model, we have to add monthly country-level emission values to the current VAR dataset.

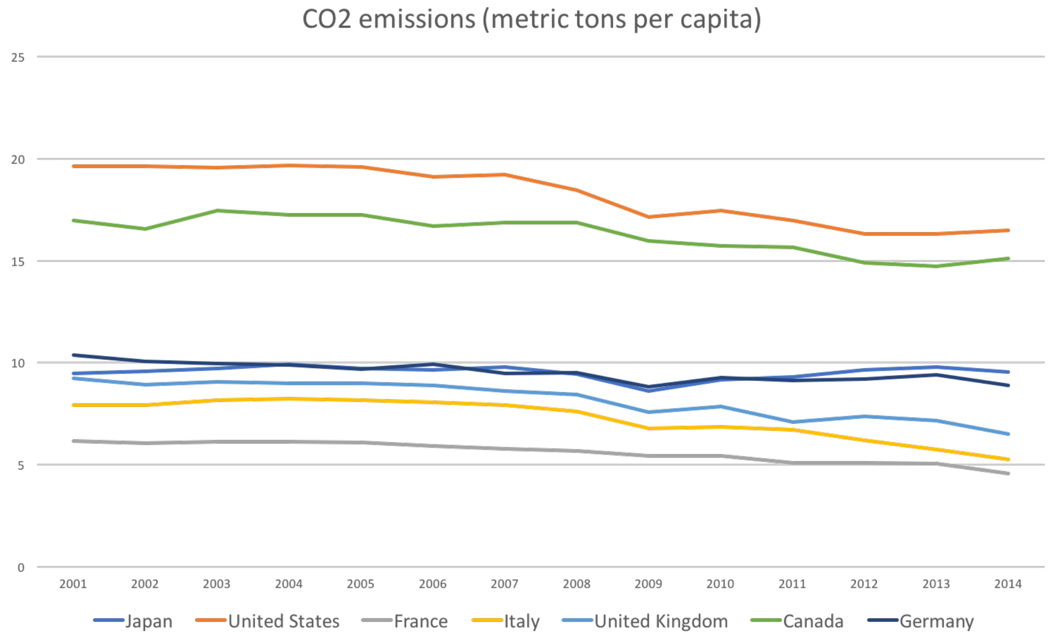


Figure 3.8: CO2 Emissions in G7 Countries

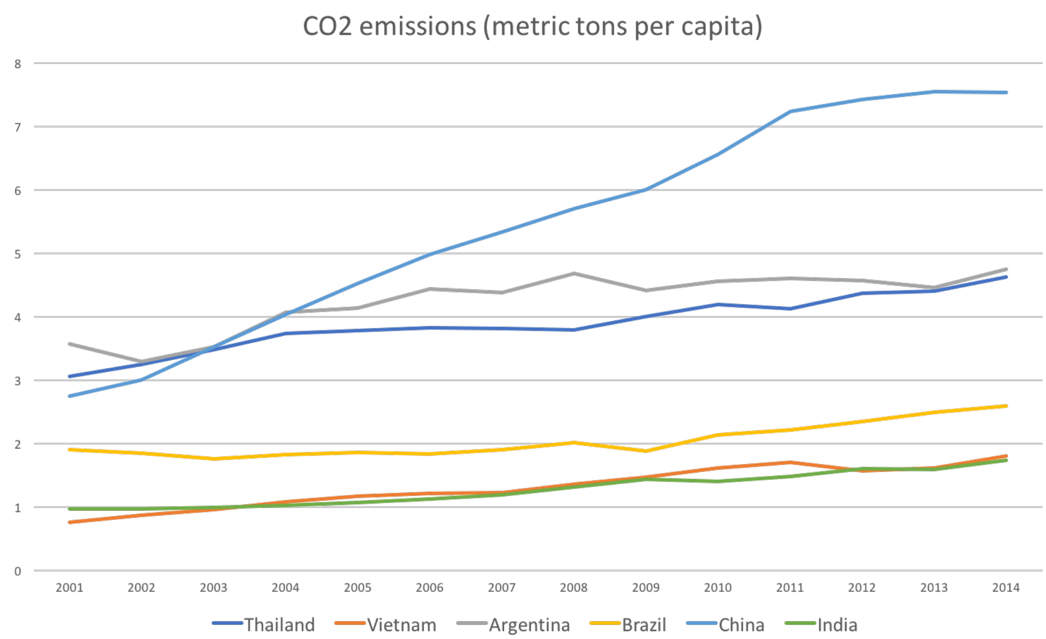


Figure 3.9: CO2 Emissions in Emerging Markets

3.4 Conclusion

To empirically demonstrate the effect of U.S. monetary policy on international trade through the channel of exchange rates, we exploit the VAR model with the dataset from China, Japan, and Korea spanning 2001-2017. Our empirical analyses show that China and Japan (large open economies) and Korea (a small open economy) respond differently to U.S. monetary policy shocks. While Korea's trade volume immediately increases following a U.S. monetary policy tightening, China's trade volume decreases. A depreciation of local currency due to a contractionary U.S. monetary policy improves the trade balance in Korea but it leads to the deteriorated trade balance in China. The depreciation of local currency in China may limit the total credits which need to be financed to sustain the global value chain and, as a result, lower total trade volume and worsen trade balance. The heterogeneous responses may be attributed to the different extent of GVC participation between two countries. In a similar setting, we will analyze how emission levels depend on monetary policy via the channel of GVC in our future work.

Bibliography

Aizenman, Joshua, Menzie D Chinn, and Hiro Ito, "Balance sheet effects on monetary and financial spillovers: The East Asian crisis plus 20," *Journal of International Money and Finance*, 2017, *74*, 258–282.

Anzuini, Alessio and Fabio Fornari, "Macroeconomic determinants of carry trade activity," *Review of International Economics*, 2012, *20* (3), 468–488.

Arize, Augustine C, John Malindretos, and Emmanuel U Igwe, "Do exchange rate changes improve the trade balance: An asymmetric nonlinear cointegration approach," *International Review of Economics & Finance*, 2017, *49*, 313–326.

Bahmani-Oskooee, Mohsen and Artatrana Ratha, "The J-curve: a literature review," *Applied Economics*, 2004, *36* (13), 1377–1398.

Banerjee, Ryan, Michael B Devereux, and Giovanni Lombardo, "Self-oriented monetary policy, global financial markets and excess volatility of international capital flows," *Journal of International Money and Finance*, 2016, *68*, 275–297.

Bauer, Michael D and Christopher J Neely, "International channels of the Fed's unconventional monetary policy," *Journal of International Money and Finance*, 2014, *44*, 24–46.

Bems, Rudolfs, Robert C Johnson, and Kei-Mu Yi, "Vertical linkages and the collapse of global trade," *American Economic Review*, 2011, *101* (3), 308–12.

Bjørnland, Hilde C and Kai Leitemo, "Identifying the interdependence between US monetary policy and the stock market," *Journal of Monetary Economics*, 2009, *56* (2), 275–282.

Borio, Claudio and Haibin Zhu, "Capital regulation, risk-taking and monetary policy: a missing link in the transmission mechanism?," *Journal of Financial Stability*, 2012, *8* (4), 236–251.

Boudoukh, Jacob, Matthew Richardson, and Robert F Whitelaw, "New evidence on the forward premium puzzle," *Journal of Financial and Quantitative Analysis*, 2016, *51* (3), 875–897.

Bowman, David, Juan M Londono, and Horacio Sapriza, "US unconventional monetary policy and transmission to emerging market economies," *Journal of International Money and Finance*, 2015, *55*, 27–59.

Bräuning, Falk and Victoria Ivashina, "US monetary policy and emerging market credit cycles," *Journal of Monetary Economics*, 2019.

Bruno, Valentina and Hyun Song Shin, "Capital flows and the risk-taking channel of monetary policy," *Journal of Monetary Economics*, 2015, *71*, 119–132.

_ , Se-Jik Kim, and Hyun Shin, "Exchange rates and the working capital channel of trade fluctuations," *AEA Papers and Proceedings*, 2018, *108*, 531–36.

Burger, John D, Francis E Warnock, and Veronica Cacdac Warnock, "The Effects of US Monetary Policy on Emerging Market Economies' Sovereign and Corporate Bond Markets," Technical Report, National Bureau of Economic Research 2017.

Caceres, Carlos, Mr Yan Carriere-Swallow, Ishak Demir, and Bertrand Gruss, *US monetary policy normalization and global interest rates*, International Monetary Fund, 2016.

Campbell, John Y, Carolin Pflueger, and Luis M Viceira, "Monetary policy drivers of bond and equity risks," Technical Report, National Bureau of Economic Research 2014.

Chen, Qianying, Andrew Filardo, Dong He, and Feng Zhu, "Financial crisis, US unconventional monetary policy and international spillovers," *Journal of International Money and Finance*, 2016, *67*, 62–81.

Connolly, Robert, Chris Stivers, and Licheng Sun, "Stock market uncertainty and the stock-bond return relation," *Journal of Financial and Quantitative Analysis*, 2005, *40* (1), 161–194.

Dornbusch, Rudiger, "Expectations and exchange rate dynamics," *Journal of Political Economy*, 1976, pp. 1161–1176.

Edwards, Sebastian, "Capital controls, capital flow contractions, and macroeconomic vulnerability," *Journal of International Money and Finance*, 2007, *26* (5), 814–840.

Eichenbaum, Martin and Charles L Evans, "Some empirical evidence on the effects of shocks to monetary policy on exchange rates," *Quarterly Journal of Economics*, 1995, *110* (4), 975–1009.

Gagnon, Joseph, Tamim Bayoumi, Juan M Londono, Christian Saborowski, and Horacio Sapriza, "Unconventional Monetary and Exchange Rate Policies," 2017.

Galí, Jordi, "Monetary policy and rational asset price bubbles," *American Economic Review*, 2014, *104* (3), 721–52.

_ **and Luca Gambetti**, "The effects of monetary policy on stock market bubbles: Some evidence," *American Economic Journal: Macroeconomics*, 2015, *7* (1), 233–57.

Gertler, Mark and Peter Karadi, "Monetary policy surprises, credit costs, and economic activity," *American Economic Journal: Macroeconomics*, 2015, *7* (1), 44–76.

Gilchrist, S, E Zakrajsek, and VZ Yue, "The Response of Sovereign Bond Yields to US Monetary Policy," *Boston University and Federal Reserve Board mimeo*, 2015.

Gilchrist, Simon, David López-Salido, and Egon Zakrajšek, "Monetary policy and real borrowing costs at the zero lower bound," *American Economic Journal: Macroeconomics*, 2015, *7* (1), 77–109.

_ **, Vivian Z Yue, and Egon Zakrajsek**, "US Monetary Policy and International Bond Markets," 2018.

Gürkaynak, Refet S, Brian Sack, and Eric Swanson, "The sensitivity of long-term interest rates to economic news: evidence and implications for macroeconomic models," *American Economic Review*, 2005, *95* (1), 425–436.

Hausman, Joshua and Jon Wongswan, "Global asset prices and FOMC announcements," *Journal of International Money and Finance*, 2011, *30* (3), 547–571.

Kurov, Alexander and Raluca Stan, "Monetary policy uncertainty and the market reaction to macroeconomic news," *Journal of Banking & Finance*, 2018, *86*, 127–142.

Kuttner, Kenneth N, "Monetary policy surprises and interest rates: Evidence from the Fed funds futures market," *Journal of Monetary Economics*, 2001, *47* (3), 523–544.

Leigh, Mr Daniel, Weicheng Lian, Marcos Poplawski-Ribeiro, Rachel Szymanski, Viktor Tsyrennikov, and Hong Yang, *Exchange rates and trade: A disconnect?*, International Monetary Fund, 2017.

McCauley, Robert N, Patrick McGuire, and Vladyslav Sushko, "Global dollar credit: links to US monetary policy and leverage," *Economic Policy*, 2015, *30* (82), 187–229.

Meinusch, Annette and Peter Tillmann, "The macroeconomic impact of unconventional monetary policy shocks," *Journal of Macroeconomics*, 2016, *47*, 58–67.

Morell, Joseph, "The decline in the predictive power of the US term spread: A structural interpretation," *Journal of Macroeconomics*, 2018, *55*, 314–331.

Moulton, Brent R, "Random group effects and the precision of regression estimates," *Journal of Econometrics*, 1986, *32* (3), 385–397.

Neely, Christopher J, "Unconventional monetary policy had large international effects," *Journal of Banking & Finance*, 2015, *52*, 101–111.

Pagan, Adrian, "Econometric issues in the analysis of regressions with generated regressors," *International Economic Review*, 1984, pp. 221–247.

Rey, Hélène, "International channels of transmission of monetary policy and the mundellian trilemma," *IMF Economic Review*, 2016, *64* (1), 6–35.

Rogoff, Kenneth, "The purchasing power parity puzzle," *Journal of Economic Literature*, 1996, *34* (2), 647–668.

Smets, Frank and Rafael Wouters, "Shocks and frictions in US business cycles: A Bayesian DSGE approach," *American Economic Review*, 2007, *97* (3), 586–606.

Sun, Chia-Hung and Yi-Bin Chiu, "Taiwan's trade imbalance and exchange rate revisited," *Applied Economics*, 2010, *42* (7), 917–922.

Swanson, Eric T and John C Williams, "Measuring the effect of the zero lower bound on medium-and longer-term interest rates," *American Economic Review*, 2014, *104* (10), 3154–85.

Taylor, John, "International monetary policy coordination: past, present and future," 2013.

Tillmann, Peter, "Unconventional monetary policy and the spillovers to emerging markets," *Journal of International Money and Finance*, 2016, *66*, 136–156.

Wong, Anson and Daniel Cheung, "Put All the Eggs in the Same Basket–An Empirical Examination of US Monetary Policy Stance and Asian Stock Markets Integration," *Available at SSRN 2828598*, 2016.

Wongswan, Jon, "The response of global equity indexes to US monetary policy announcements," *Journal of International Money and Finance*, 2009, *28* (2), 344–365.

Yellen, Janet L, "Revolution and evolution in central bank communications," *Speech given at the University of California Haas School of Business, Berkeley, California November*, 2012, *13* (0).

Appendix for Chapter 1: Tables

Country	CAN	CHL	COL	ZAF	AUS	NZL
	(1)	(2)	(3)	(4)	(5)	(6)
VARIABLES	GOV2	GOV2	GOV2	GOV2	GOV2	GOV2
FF	0.487**	-0.284	0.234	0.382***	0.386*	0.209
	(0.208)	(0.418)	(0.141)	(0.121)	(0.199)	(0.177)
TYF	0.0507**	0.00397	-0.0451	0.00896	0.0428**	-0.00559
	(0.0201)	(0.0397)	(0.0348)	(0.0362)	(0.0173)	(0.0125)
FF × CRISIS	0.278	0.879*	0.465**	0.981***	-0.444*	-3.372
	(0.246)	(0.514)	(0.226)	(0.260)	(0.229)	(2.981)
TYF × CRISIS	-0.0296	-0.0200	0.0152	0.0272	0.0195	-0.139
	(0.0212)	(0.0412)	(0.0375)	(0.0394)	(0.0211)	(0.183)
FF × POST	0.0672	-2.353***	1.131	3.472***	1.144**	4.443
	(0.410)	(0.354)	(2.601)	(0.476)	(0.478)	(3.047)
TYF × POST	-0.0296	-0.0292	0.0654	0.278***	0.0152	0.161
	(0.0514)	(0.0334)	(0.0885)	(0.0359)	(0.0522)	(0.189)
CRISIS	-0.00219	0.00485	0.00657	0.0187	-0.00997	-0.00981
	(0.00923)	(0.0214)	(0.0115)	(0.0231)	(0.0107)	(0.0137)
POST	-0.00552	-0.0167	0.00988	-0.00139	-0.0152	0.00120
	(0.00887)	(0.0111)	(0.0138)	(0.0218)	(0.0120)	(0.0148)
Constant	0.00515	-0.00860	-0.0155	-0.0174	0.0157*	0.00720
	(0.00781)	(0.0202)	(0.00951)	(0.0108)	(0.00826)	(0.00759)
Observations	136	75	104	88	136	91
Adjusted R-squared	0.222	-0.0270	0.0258	0.207	0.191	0.0208

NOTE: The dependent variable is daily change in 2-year (GOV2) ahead government bond yield bracketing an FOMC announcement. The entries labeled "FF" denote a 30-minute window change in the Fed Fund Futures around an FOMC announcement. The entries labeled "TYF" denote a 30-minute change in the 10-year Treasury Futures. "CRISIS" is 1 in the sample period between Nov 2008 and Dec 2015. "POST" is 1 in the sample period after Dec 2015. Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1

Table A1: Response of 2-year Government Bond Yield to U.S. Monetary Policy Surprises

Country	AUT	BEL	FRA	DEU	NLD	CHE	GRC	ITA	PRT	ESP
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
VARIABLES	GOV2	GOV2	GOV2	GOV2	GOV2	GOV2	GOV2	GOV2	GOV2	GOV2
FF	0.442**	0.448**	0.318*	0.378*	0.295*	-0.542	0.364**	0.352**	0.454**	0.360**
	(0.182)	(0.225)	(0.170)	(0.194)	(0.175)	(0.573)	(0.158)	(0.170)	(0.211)	(0.171)
TYF	0.0276	0.0348	0.0323	0.0457***	0.0334	-0.175	-0.0538**	0.0381*	0.0368*	0.0318
	(0.0227)	(0.0221)	(0.0218)	(0.0168)	(0.0218)	(0.131)	(0.0205)	(0.0222)	(0.0216)	(0.0226)
FF × CRISIS	0.725***	0.431	0.709***	0.874***	0.907***	0.981*	8.525	0.425	0.0774	0.193
	(0.219)	(0.317)	(0.195)	(0.213)	(0.207)	(0.577)	(11.95)	(0.423)	(0.610)	(0.491)
TYF × CRISIS	-0.0181	-0.0270	-0.0234	-0.0421**	-0.0276	0.157	0.309	-0.0283	0.0209	-0.0219
	(0.0260)	(0.0235)	(0.0250)	(0.0210)	(0.0263)	(0.131)	(0.281)	(0.0248)	(0.0443)	(0.0266)
FF × POST	-1.092***	-0.884***	-1.470***	-1.184***	-1.251***	0.00267	-4.236	-1.182**	-0.747	-0.949*
	(0.177)	(0.312)	(0.256)	(0.167)	(0.187)	(0.374)	(13.09)	(0.475)	(2.007)	(0.545)
TYF × POST	0.0569***	0.0483**	0.0509**	0.0573**	0.0482**	0.00464	-0.0319	0.0647	0.112	0.0446*
	(0.0194)	(0.0211)	(0.0244)	(0.0247)	(0.0232)	(0.0288)	(0.496)	(0.0498)	(0.182)	(0.0257)
CRISIS	-0.0254***	-0.0181*	-0.0201**	-0.0240***	-0.0193**	0.0344	0.628	-0.0205	0.0158	-0.0314**
	(0.00889)	(0.00928)	(0.00866)	(0.00888)	(0.00872)	(0.0447)	(0.834)	(0.0133)	(0.0303)	(0.0143)
POST	0.0219***	0.0125	0.0150**	0.0168**	0.0162**	0.00582	-0.672	0.0133	0.0248	0.0248*
	(0.00670)	(0.00815)	(0.00747)	(0.00692)	(0.00662)	(0.00715)	(0.837)	(0.0142)	(0.0550)	(0.0136)
Constant	0.00928	0.00427	0.00365	0.00975	0.00470	-0.0418	0.00467	0.00749	0.00581	0.00662
	(0.00649)	(0.00618)	(0.00646)	(0.00686)	(0.00661)	(0.0445)	(0.00992)	(0.00622)	(0.00654)	(0.00605)
Observations	134	135	136	136	135	91	62	136	136	136
Adjusted R-squared	0.252	0.175	0.197	0.267	0.219	0.157	-0.135	0.0554	-0.0169	0.0511

NOTE: The dependent variable is daily change in 2-year (GOV2) ahead government bond yield bracketing an FOMC announcement. The entries labeled "FF" denote a 30-minute window change in the Fed Fund Futures around an FOMC announcement. The entries labeled "TYF" denote a 30-minute change in the 10-year Treasury Futures. "CRISIS" is 1 in the sample period between Nov 2008 and Dec 2015. "POST" is 1 in the sample period after Dec 2015. Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1

Table A1: Response of 2-year Government Bond Yield (Continued)

Country	DNK	FIN	IRL	NOR	SWE	GBR	CZE	HUN	POL	CHN
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
VARIABLES	GOV2	GOV2	GOV2	GOV2	GOV2	GOV2	GOV2	GOV2	GOV2	GOV2
FF	0.377**	0.187	0.0411	-0.0305	0.347*	0.236	0.121	0.525	0.204**	-0.136
	(0.164)	(0.130)	(0.0740)	(0.0265)	(0.204)	(0.198)	(0.0874)	(0.560)	(0.101)	(0.0974)
TYF	0.0250	0.0285	0.0295	-0.0478	-0.0597**	0.00751	0.0185	0.499	0.00316	0.00638
	(0.0205)	(0.0193)	(0.0321)	(0.0296)	(0.0253)	(0.0171)	(0.0218)	(0.371)	(0.0396)	(0.0191)
FF × CRISIS	0.571***	0.783***	0.407	0.398	1.011***	0.639***	-0.0484	0.277	-0.0585	0.445***
	(0.197)	(0.162)	(3.780)	(1.161)	(0.270)	(0.221)	(0.162)	(0.651)	(0.156)	(0.132)
TYF × CRISIS	-0.0112	-0.0181	0.0183	0.0759**	0.0567*	0.00331	-0.0197	-0.450	0.0169	-0.00259
	(0.0248)	(0.0228)	(0.0713)	(0.0342)	(0.0299)	(0.0189)	(0.0229)	(0.371)	(0.0410)	(0.0195)
FF × POST	0.114	-0.782***	-0.489	2.714	-0.811**	-0.846**	-0.597**	-0.164	-0.170	0.993***
	(0.285)	(0.199)	(3.783)	(5.758)	(0.341)	(0.331)	(0.297)	(1.191)	(0.373)	(0.137)
TYF × POST	-0.0161	0.0435*	-0.00308	-0.0812	0.00835	0.102**	0.0556	0.0760	0.0694**	-0.0997***
	(0.0297)	(0.0256)	(0.0656)	(0.0795)	(0.0362)	(0.0427)	(0.0346)	(0.120)	(0.0276)	(0.0184)
CRISIS	-0.0171**	-0.0143*	-0.0130	0.0150	-0.0231**	-0.0130	-0.0249**	0.0585	0.00565	0.0105
	(0.00842)	(0.00807)	(0.0314)	(0.0159)	(0.0115)	(0.00924)	(0.0118)	(0.0417)	(0.0105)	(0.00941)
POST	0.0216***	0.0181**	0.0160	0.0188	0.0121	0.0189**	0.0136	-0.00577	-0.00615	-0.0115**
	(0.00745)	(0.00695)	(0.0304)	(0.0207)	(0.00876)	(0.00894)	(0.00915)	(0.0207)	(0.00745)	(0.00558)
Constant	0.00315	-8.65e-05	-0.000915	-0.0191***	0.0122	0.00326	0.0129	-0.0614	0.000440	-0.00466
	(0.00601)	(0.00557)	(0.00890)	(0.00660)	(0.00920)	(0.00686)	(0.0101)	(0.0391)	(0.00850)	(0.00815)
Observations	136	125	101	39	92	136	91	91	136	81
Adjusted R-squared	0.200	0.154	-0.0701	-0.116	0.178	0.104	-0.0265	0.181	-0.00991	-0.0130

NOTE: The dependent variable is daily change in 2-year (GOV2) ahead government bond yield bracketing an FOMC announcement. The entries labeled "FF" denote a 30-minute window change in the Fed Fund Futures around an FOMC announcement. The entries labeled "TYF" denote a 30-minute change in the 10-year Treasury Futures. "CRISIS" is 1 in the sample period between Nov 2008 and Dec 2015. "POST" is 1 in the sample period after Dec 2015. Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1

Table A1: Response of 2-year Government Bond Yield (Continued)

Country	HKG	JPN	KOR	TWN	IND	IDN	MYS	SGP	THA	TUR
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
VARIABLES	GOV2	GOV2	GOV2	GOV2	GOV2	GOV2	GOV2	GOV2	GOV2	GOV2
FF	0.727***	0.0146	0.138	0.206*	0.0232	1.508	0.0564	0.0325	0.249**	1.079*
	(0.208)	(0.0350)	(0.123)	(0.115)	(0.0822)	(1.648)	(0.0636)	(0.0955)	(0.103)	(0.599)
TYF	0.0771*	0.00374	0.0265**	0.00868	0.0240	0.232	0.00925	0.0661	0.0210	0.511**
	(0.0426)	(0.00493)	(0.0124)	(0.0116)	(0.0229)	(0.219)	(0.0110)	(0.0441)	(0.0137)	(0.197)
FF × CRISIS	-0.117	0.238***	0.543**	0.686	-0.717*	3.711**	0.535***	0.157	1.067***	3.169***
	(0.230)	(0.0566)	(0.248)	(0.637)	(0.421)	(1.740)	(0.0923)	(0.103)	(0.260)	(0.837)
TYF × CRISIS	-0.0433	-0.00349	-0.0133	-0.0642	0.00293	-0.218	-0.000740	-0.0565	-0.0155	-0.464**
	(0.0434)	(0.00551)	(0.0273)	(0.0463)	(0.0430)	(0.221)	(0.0123)	(0.0443)	(0.0174)	(0.199)
FF × POST	2.589***	0.343***	0.191	-0.279	0.785	-5.193***	0.351	1.919***	-0.413	2.821***
	(0.746)	(0.119)	(0.354)	(0.636)	(0.555)	(0.831)	(0.239)	(0.391)	(0.482)	(0.680)
TYF × POST	-0.0244	0.00159	-0.0547	0.0789	0.0398	0.146	-0.00860	0.0534	0.0439	-0.0116
	(0.120)	(0.0111)	(0.0345)	(0.0519)	(0.0640)	(0.104)	(0.0311)	(0.0687)	(0.0575)	(0.0595)
CRISIS	-0.00338	-0.00214	0.0147	-0.00151	-0.0103	0.0537	0.00638	-0.0105	-0.00700	-0.0240
	(0.0105)	(0.00221)	(0.0112)	(0.0149)	(0.0253)	(0.0889)	(0.00728)	(0.00803)	(0.0112)	(0.0802)
POST	-0.00866	0.00149	-0.00971	0.0144	0.0218	0.0366	0.00522	0.000202	-0.00767	0.0776**
	(0.0156)	(0.00246)	(0.00931)	(0.0125)	(0.0244)	(0.0250)	(0.00634)	(0.0119)	(0.00926)	(0.0360)
Constant	0.00241	0.000700	-0.0129	-0.00711	-0.0105	-0.0823	-0.00760	0.0130*	0.00927	-0.0148
	(0.00942)	(0.00194)	(0.00795)	(0.00914)	(0.0111)	(0.0871)	(0.00644)	(0.00753)	(0.00847)	(0.0726)
Observations	130	136	120	96	110	115	119	91	118	89
Adjusted R-squared	0.290	0.0856	0.0342	0.00727	-0.0614	0.0535	0.0681	0.190	0.145	0.0817

NOTE: The dependent variable is daily change in 2-year (GOV2) ahead government bond yield bracketing an FOMC announcement. The entries labeled "FF" denote a 30-minute window change in the Fed Fund Futures around an FOMC announcement. The entries labeled "TYF" denote a 30-minute change in the 10-year Treasury Futures. "CRISIS" is 1 in the sample period between Nov 2008 and Dec 2015. "POST" is 1 in the sample period after Dec 2015. Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1

Table A1: Response of 2-year Government Bond Yield (Continued)

Country	CAN	BRA	CHL	COL	ZAF	AUS	NZL
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
VARIABLES	GOV10	GOV10	GOV10	GOV10	GOV10	GOV10	GOV10
FF	0.207	0.872***	-2.088***	2.459***	0.184	0.360	0.249
	(0.127)	(0.135)	(0.218)	(0.269)	(0.237)	(0.234)	(0.170)
TYF	0.0511***	0.429***	0.186***	-0.193**	0.0325	0.0783***	0.00475
	(0.0161)	(0.0869)	(0.0520)	(0.0885)	(0.0502)	(0.0178)	(0.0188)
FF × CRISIS	0.0781	-6.277	2.460***	-1.480***	0.924**	-0.393	-0.589***
	(0.167)	(3.771)	(0.769)	(0.330)	(0.443)	(0.269)	(0.214)
TYF × CRISIS	0.0204	-0.174	-0.160	0.179**	0.0105	0.0202	0.0508**
	(0.0182)	(0.114)	(0.111)	(0.0902)	(0.0601)	(0.0205)	(0.0213)
FF × POST	0.100	9.403**	-0.0260	0.746	4.455***	1.702**	2.221**
	(0.474)	(4.484)	(0.747)	(1.357)	(0.902)	(0.839)	(0.923)
TYF × POST	-0.0827	-0.416*	-0.0258	-0.00953	0.314***	-0.0782	-0.0763
	(0.0670)	(0.241)	(0.101)	(0.0981)	(0.0751)	(0.0845)	(0.0884)
CRISIS	0.00103	0.105***	-0.00398	0.0300	-0.00411	-0.0114	-0.0109
	(0.00887)	(0.0304)	(0.0213)	(0.0453)	(0.0176)	(0.0128)	(0.0114)
POST	-0.0233*	-0.0471	0.0164	0.0160	-0.000888	-0.0282	-0.0187
	(0.0126)	(0.0393)	(0.0134)	(0.0195)	(0.0191)	(0.0185)	(0.0185)
Constant	0.00257	-0.0641**	-0.00825	-0.0409	0.00158	0.0154*	0.0116
	(0.00564)	(0.0243)	(0.0173)	(0.0442)	(0.0108)	(0.00882)	(0.00736)
Observations	136	57	40	91	126	136	136
Adjusted R-squared	0.269	0.229	0.304	0.122	0.169	0.267	0.0724

NOTE: The dependent variable is daily change in 10-year (GOV10) ahead government bond yield bracketing an FOMC announcement. The entries labeled "FF" denote a 30-minute window change in the Fed Fund Futures around an FOMC announcement. The entries labeled "TYF" denote a 30-minute change in the 10-year Treasury Futures. "CRISIS" is 1 in the sample period between Nov 2008 and Dec 2015. "POST" is 1 in the sample period after Dec 2015. Robust standard errors in parentheses.
*** p<0.01, ** p<0.05, * p<0.1

Table A2: Response of 10-year Government Bond Yield to U.S. Monetary Policy Surprises

Country	AUT	BEL	FRA	DEU	NLD	CHE	GRC	ITA	PRT	ESP
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
VARIABLES	GOV10	GOV10	GOV10	GOV10	GOV10	GOV10	GOV10	GOV10	GOV10	GOV10
FF	0.304	0.215	0.288	0.283	0.262	-0.123	0.0623	0.270	0.200	0.265
	(0.194)	(0.188)	(0.200)	(0.196)	(0.179)	(0.177)	(0.199)	(0.185)	(0.194)	(0.190)
TYF	0.0273	0.0343	0.0331	0.0319	0.0329	-0.0407*	-0.0693**	0.0310	0.0269	0.0333
	(0.0208)	(0.0213)	(0.0215)	(0.0216)	(0.0209)	(0.0211)	(0.0307)	(0.0215)	(0.0217)	(0.0221)
FF × CRISIS	0.324	0.283	0.273	0.207	0.276	0.362	-0.499	0.306	0.171	0.208
	(0.244)	(0.221)	(0.238)	(0.259)	(0.232)	(0.222)	(1.065)	(0.330)	(0.355)	(0.347)
TYF × CRISIS	0.0130	0.00839	0.0108	0.0211	0.0155	0.0735***	0.0775	0.0139	0.0327	0.0191
	(0.0228)	(0.0232)	(0.0233)	(0.0243)	(0.0235)	(0.0234)	(0.0618)	(0.0242)	(0.0311)	(0.0269)
FF × POST	0.424	0.851	0.852	0.878	0.682	0.444	1.938	0.268	1.103	0.964
	(0.636)	(0.657)	(0.630)	(0.757)	(0.707)	(0.583)	(1.646)	(0.672)	(1.046)	(0.609)
TYF × POST	0.164***	0.0973**	0.101**	0.139***	0.129***	0.0402	0.128	0.0935*	0.0893	0.0430
	(0.0528)	(0.0427)	(0.0433)	(0.0439)	(0.0429)	(0.0321)	(0.151)	(0.0539)	(0.0952)	(0.0499)
CRISIS	-0.0242**	-0.0249**	-0.0234**	-0.0240**	-0.0253**	0.000114	-0.0557	-0.0172	-0.0207	-0.0144
	(0.0101)	(0.0105)	(0.0102)	(0.0111)	(0.0108)	(0.0112)	(0.0470)	(0.0137)	(0.0190)	(0.0149)
POST	0.0311**	0.0165	0.0171	0.0202*	0.0214*	0.00819	0.0668	0.00997	0.0337	0.0122
	(0.0154)	(0.0108)	(0.0106)	(0.0122)	(0.0115)	(0.00891)	(0.0521)	(0.0156)	(0.0291)	(0.0159)
Constant	0.00684	0.00778	0.00702	0.00790	0.00770	-0.00393	-0.000223	0.00789	0.00570	0.00717
	(0.00698)	(0.00701)	(0.00713)	(0.00685)	(0.00683)	(0.00940)	(0.0133)	(0.00709)	(0.00784)	(0.00700)
Observations	135	136	136	136	136	90	91	136	131	136
Adjusted R-squared	0.195	0.174	0.195	0.194	0.191	0.0790	-0.0826	0.0889	0.0232	0.0736

NOTE: The dependent variable is daily change in 10-year (GOV10) ahead government bond yield bracketing an FOMC announcement. The entries labeled "FF" denote a 30-minute window change in the Fed Fund Futures around an FOMC announcement. The entries labeled "TYF" denote a 30-minute change in the 10-year Treasury Futures. "CRISIS" is 1 in the sample period between Nov 2008 and Dec 2015. "POST" is 1 in the sample period after Dec 2015. Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1

Table A2: Response of 10-year Government Bond Yield (Continued)

Country	DNK	FIN	IRL	NOR	SWE	GBR	CZE	HUN	POL	CHN
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
VARIABLES	GOV10	GOV10	GOV10	GOV10	GOV10	GOV10	GOV10	GOV10	GOV10	GOV10
FF	0.308	0.307	0.434***	0.166	0.129	0.283	0.0879	0.411	0.232**	-0.0441
	(0.197)	(0.190)	(0.163)	(0.239)	(0.212)	(0.180)	(0.152)	(0.416)	(0.103)	(0.0654)
TYF	0.0260	0.0325	0.0322	-0.101**	-0.0599**	0.0103	-0.153	0.417	-0.0249	0.0212
	(0.0209)	(0.0213)	(0.0219)	(0.0470)	(0.0250)	(0.0228)	(0.114)	(0.307)	(0.0425)	(0.0252)
FF × CRISIS	0.747	0.239	0.285	0.904***	1.100***	0.502*	0.452**	0.892	0.108	0.382***
	(1.385)	(0.244)	(0.221)	(0.278)	(0.252)	(0.264)	(0.188)	(0.727)	(0.191)	(0.114)
TYF × CRISIS	0.0206	0.0148	-0.00454	0.140***	0.0978***	0.0313	0.159	-0.518	0.0503	-0.0176
	(0.0236)	(0.0239)	(0.0245)	(0.0484)	(0.0277)	(0.0296)	(0.114)	(0.315)	(0.0444)	(0.0261)
FF × POST	0.631	0.703	0.603	1.020***	0.637	-0.00735	-1.090***	0.214	0.0782	-0.642
	(1.504)	(0.675)	(0.598)	(0.382)	(0.723)	(0.911)	(0.332)	(1.187)	(1.002)	(0.400)
TYF × POST	0.110**	0.125***	0.113**	0.124***	0.0534	0.221***	0.127***	0.203*	0.209**	0.133**
	(0.0426)	(0.0380)	(0.0491)	(0.0297)	(0.0560)	(0.0610)	(0.0272)	(0.106)	(0.0938)	(0.0654)
CRISIS	-0.0224*	-0.0249**	-0.00471	-0.00755	-0.0157	-0.0122	-0.0286*	0.0696*	0.0114	-0.00647
	(0.0114)	(0.0108)	(0.0136)	(0.0175)	(0.0119)	(0.0130)	(0.0156)	(0.0388)	(0.0131)	(0.00776)
POST	0.0180	0.0178	0.00753	0.0143	0.0142	0.0122	0.00573	-0.0112	-0.0198	0.0154
	(0.0118)	(0.0111)	(0.0157)	(0.0101)	(0.0122)	(0.0146)	(0.00855)	(0.0226)	(0.0152)	(0.0125)
Constant	0.00870	0.00776	0.00410	0.00338	0.00712	0.00739	0.0247*	-0.0539	0.000158	0.00102
	(0.00690)	(0.00692)	(0.00708)	(0.0151)	(0.00897)	(0.00766)	(0.0143)	(0.0339)	(0.00785)	(0.00620)
Observations	128	136	119	91	92	136	91	91	128	83
Adjusted R-squared	0.147	0.195	0.111	0.212	0.246	0.112	0.161	0.147	0.0427	0.0608

NOTE: The dependent variable is daily change in 10-year (GOV10) ahead government bond yield bracketing an FOMC announcement. The entries labeled "FF" denote a 30-minute window change in the Fed Fund Futures around an FOMC announcement. The entries labeled "TYF" denote a 30-minute change in the 10-year Treasury Futures. "CRISIS" is 1 in the sample period between Nov 2008 and Dec 2015. "POST" is 1 in the sample period after Dec 2015. Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1

Table A2: Response of 10-year Government Bond Yield (Continued)

Country	HKG	JPN	KOR	TWN	IND	IDN	MYS	PHL	SGP	THA
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
VARIABLES	GOV10	GOV10	GOV10	GOV10	GOV10	GOV10	GOV10	GOV10	GOV10	GOV10
FF	0.546***	-0.125*	-0.00882	0.139	0.320***	-0.0530	0.195**	0.0566	0.238***	0.380
	(0.176)	(0.0698)	(0.218)	(0.0847)	(0.109)	(0.398)	(0.0952)	(0.182)	(0.0571)	(0.247)
TYF	0.0770***	0.00952	0.0573**	0.0189	0.0448***	0.0709	-0.00974	0.0267	0.0288	0.0167
	(0.0157)	(0.0108)	(0.0243)	(0.0125)	(0.0156)	(0.0541)	(0.0305)	(0.0190)	(0.0258)	(0.0217)
FF × CRISIS	0.0131	0.467***	0.125	0.359**	0.565***	7.943***	0.412***	-0.304	-0.370*	0.668**
	(0.219)	(0.0792)	(0.248)	(0.137)	(0.160)	(0.798)	(0.114)	(0.684)	(0.191)	(0.278)
TYF × CRISIS	-0.00440	0.00369	0.00552	-0.0151	-0.0363*	0.0469	0.0318	0.0246	0.0238	0.0481*
	(0.0199)	(0.0112)	(0.0275)	(0.0151)	(0.0185)	(0.0645)	(0.0308)	(0.0235)	(0.0332)	(0.0260)
FF × POST	1.375*	-1.270***	1.106**	2.708	0.282	-4.936***	1.353***	19.48*	3.813***	0.911*
	(0.704)	(0.228)	(0.523)	(5.582)	(0.497)	(1.546)	(0.401)	(11.10)	(0.554)	(0.464)
TYF × POST	-0.0696	0.00304	-0.0725	-0.0529	0.113*	0.0618	0.0473	0.0721	0.0435	-0.0800
	(0.110)	(0.0259)	(0.0554)	(0.0659)	(0.0638)	(0.155)	(0.0528)	(0.167)	(0.0814)	(0.0588)
CRISIS	0.00613	-0.00368	0.0261*	-0.00151	-9.40e-05	0.00148	0.0116	0.0218*	-0.0335**	-0.00451
	(0.0106)	(0.00474)	(0.0134)	(0.00700)	(0.0108)	(0.0311)	(0.0105)	(0.0128)	(0.0132)	(0.0141)
POST	-0.000567	0.00278	-0.0216*	-0.00662	0.0193	0.00515	0.00480	-0.00715	-0.00391	-0.00839
	(0.0237)	(0.00527)	(0.0116)	(0.0193)	(0.0130)	(0.0278)	(0.0106)	(0.0237)	(0.0139)	(0.0122)
Constant	-0.00406	-0.000615	-0.0237**	0.000875	-0.00678	-0.0225	-0.00891	-0.0171	0.0269**	0.00210
	(0.00724)	(0.00400)	(0.0116)	(0.00566)	(0.00868)	(0.0233)	(0.00965)	(0.0118)	(0.0112)	(0.0115)
Observations	132	136	119	87	134	115	99	59	91	119
Adjusted R-squared	0.313	0.0969	0.138	0.0614	0.115	0.502	0.211	-0.0427	0.230	0.237

NOTE: The dependent variable is daily change in 10-year (GOV10) ahead government bond yield bracketing an FOMC announcement. The entries labeled "FF" denote a 30-minute window change in the Fed Fund Futures around an FOMC announcement. The entries labeled "TYF" denote a 30-minute change in the 10-year Treasury Futures. "CRISIS" is 1 in the sample period between Nov 2008 and Dec 2015. "POST" is 1 in the sample period after Dec 2015. Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1

Table A2: Response of 10-year Government Bond Yield (Continued)

Appendix for Chapter 2: Tables

	Augmented Dickey-Fuller test statistic		
Variable	Level	First Difference	Integration Order
Industrial Production	-0.967595	-12.96751	1
CPI	-1.510499	-10.93257	1
2-year Government Bond Yield	-1.894236	-13.5332	1
Stock Price Index	-2.316984	-9.40504	1
Term Spread	-1.871852	-13.67222	1

	Phillips-Perron test statistic		
Variable	Level	First Difference	Integration Order
Industrial Production	-1.425869	-13.24182	1
CPI	-1.564921	-11.69529	1
2-year Government Bond Yield	-1.915519	-13.62677	1
Stock Price Index	-1.88556	-9.468135	1
Term Spread	-1.895979	-13.67655	1

Table A3: Unit Root Tests (Canada)

	Augmented Dickey-Fuller test statistic		
Variable	Level	First Difference	Integration Order
Industrial Production	-1.623387	-19.5888	1
CPI	-3.039153	-11.41695	1
2-year Government Bond Yield	-1.077354	-11.61522	1
Stock Price Index	-2.255987	-6.494133	1
Term Spread	-1.920477	-14.0399	1

	Phillips-Perron test statistic		
Variable	Level	First Difference	Integration Order
Industrial Production	-1.934288	-18.86119	1
CPI	-2.624453	-11.68228	1
2-year Government Bond Yield	-1.258616	-11.71424	1
Stock Price Index	-1.937741	-10.16439	1
Term Spread	-2.099139	-14.05242	1

Table A4: Unit Root Tests (France)

	Augmented Dickey-Fuller test statistic		
Variable	Level	First Difference	Integration Order
Industrial Production	-2.184934	-5.400325	1
CPI	-0.617013	-13.50081	1
2-year Government Bond Yield	-1.069051	-6.646733	1
Stock Price Index	-1.149966	-9.331185	1
Term Spread	-2.000175	-13.64221	1

	Phillips-Perron test statistic		
Variable	Level	First Difference	Integration Order
Industrial Production	-1.457151	-15.07688	1
CPI	-0.617013	-13.50081	1
2-year Government Bond Yield	-1.257269	-10.27802	1
Stock Price Index	-1.069054	-10.54221	1
Term Spread	-2.159308	-13.65815	1

Table A5: Unit Root Tests (Germany)

	Augmented Dickey-Fuller test statistic		
Variable	Level	First Difference	Integration Order
Industrial Production	-1.357344	-5.619542	1
CPI	-2.482174	-6.424151	1
2-year Government Bond Yield	-1.703646	-10.26752	1
Stock Price Index	-1.838098	-6.55273	1
Term Spread	-1.77095	-10.50174	1

	Phillips-Perron test statistic		
Variable	Level	First Difference	Integration Order
Industrial Production	-1.413058	-16.28687	1
CPI	-2.364445	-11.36525	1
2-year Government Bond Yield	-1.596656	-13.53129	1
Stock Price Index	-1.733591	-10.81947	1
Term Spread	-2.619702	-16.54747	1

Table A6: Unit Root Tests (Italy)

	Augmented Dickey-Fuller test statistic		
Variable	Level	First Difference	Integration Order
Industrial Production	-2.192049	-12.87343	1
CPI	-0.472362	-11.72608	1
2-year Government Bond Yield	-1.167098	-13.223	1
Stock Price Index	-1.524861	-12.04642	1
Term Spread	-1.375837	-14.97879	1

	Phillips-Perron test statistic		
Variable	Level	First Difference	Integration Order
Industrial Production	-2.561566	-12.94605	1
CPI	-0.834847	-11.76764	1
2-year Government Bond Yield	-1.273104	-13.22271	1
Stock Price Index	-1.254977	-12.02338	1
Term Spread	-1.305924	-14.96924	1

Table A7: Unit Root Tests (Japan)

	Augmented Dickey-Fuller test statistic		
Variable	Level	First Difference	Integration Order
Industrial Production	-1.801251	-16.39846	1
CPI	0.128398	-6.742141	1
2-year Government Bond Yield	-1.148681	-10.59575	1
Stock Price Index	-2.291044	-12.04642	1
Term Spread	-1.616152	-13.27298	1

	Phillips-Perron test statistic		
Variable	Level	First Difference	Integration Order
Industrial Production	-1.706336	-16.49029	1
CPI	0.340084	-10.73248	1
2-year Government Bond Yield	-1.233328	-10.59575	1
Stock Price Index	-2.09238	-9.99597	1
Term Spread	-1.680618	-13.27053	1

Table A8: Unit Root Tests (U.K.)

	Augmented Dickey-Fuller test statistic		
Variable	Level	First Difference	Integration Order
Industrial Production	-2.624172	-3.646727	1
CPI	-1.326175	-9.069811	1
2-year Government Bond Yield	-1.348602	-9.412905	1
Stock Price Index	-0.441169	-11.73452	1
Term Spread	-1.873345	-6.317636	1

	Phillips-Perron test statistic		
Variable	Level	First Difference	Integration Order
Industrial Production	-1.713878	-12.56792	1
CPI	-1.355962	-8.090237	1
2-year Government Bond Yield	-1.54637	-9.408729	1
Stock Price Index	-0.257682	-11.87935	1
Term Spread	-1.629892	-9.667679	1

Table A9: Unit Root Tests (U.S.)

Unrestricted Cointegration Rank Test (Trace)

Hypothesized No. of CE(s)	Eigenvalue	Trace Statistic	0.05 Critical Value	Prob.**
None *	0.358392	168.5108	95.75366	0.0000
At most 1 *	0.276335	89.07455	69.81889	0.0007
At most 2	0.080121	31.18127	47.85613	0.6567
At most 3	0.049455	16.23232	29.79707	0.6959
At most 4	0.037189	7.153557	15.49471	0.5599
At most 5	0.002063	0.369731	3.841466	0.5431

Trace test indicates 2 cointegrating eqn(s) at the 0.05 level
* denotes rejection of the hypothesis at the 0.05 level
**MacKinnon-Haug-Michelis (1999) p-values

Unrestricted Cointegration Rank Test (Maximum Eigenvalue)

Hypothesized No. of CE(s)	Eigenvalue	Max-Eigen Statistic	0.05 Critical Value	Prob.**
None *	0.358392	79.43627	40.07757	0.0000
At most 1 *	0.276335	57.89327	33.87687	0.0000
At most 2	0.080121	14.94895	27.58434	0.7523
At most 3	0.049455	9.078767	21.13162	0.8260
At most 4	0.037189	6.783826	14.26460	0.5150
At most 5	0.002063	0.369731	3.841466	0.5431

Max-eigenvalue test indicates 2 cointegrating eqn(s) at the 0.05 level
* denotes rejection of the hypothesis at the 0.05 level
**MacKinnon-Haug-Michelis (1999) p-values

Table A10: Johansen's Cointegration Test (Germany)

Unrestricted Cointegration Rank Test (Trace)

Hypothesized No. of CE(s)	Eigenvalue	Trace Statistic	0.05 Critical Value	Prob.**
None *	0.278045	144.4518	95.75366	0.0000
At most 1 *	0.229567	86.13479	69.81889	0.0015
At most 2	0.106045	39.45124	47.85613	0.2427
At most 3	0.059361	19.38540	29.79707	0.4655
At most 4	0.026920	8.431394	15.49471	0.4205
At most 5	0.019619	3.546726	3.841466	0.0597

Trace test indicates 2 cointegrating eqn(s) at the 0.05 level
* denotes rejection of the hypothesis at the 0.05 level
**MacKinnon-Haug-Michelis (1999) p-values

Unrestricted Cointegration Rank Test (Maximum Eigenvalue)

Hypothesized No. of CE(s)	Eigenvalue	Max-Eigen Statistic	0.05 Critical Value	Prob.**
None *	0.278045	58.31698	40.07757	0.0002
At most 1 *	0.229567	46.68356	33.87687	0.0009
At most 2	0.106045	20.06584	27.58434	0.3366
At most 3	0.059361	10.95401	21.13162	0.6519
At most 4	0.026920	4.884667	14.26460	0.7563
At most 5	0.019619	3.546726	3.841466	0.0597

Max-eigenvalue test indicates 2 cointegrating eqn(s) at the 0.05 level
* denotes rejection of the hypothesis at the 0.05 level
**MacKinnon-Haug-Michelis (1999) p-values

Table A11: Johansen's Cointegration Test (France)

Unrestricted Cointegration Rank Test (Trace)

Hypothesized No. of CE(s)	Eigenvalue	Trace Statistic	0.05 Critical Value	Prob.**
None *	0.261502	116.1986	95.75366	0.0010
At most 1	0.157388	61.93704	69.81889	0.1807
At most 2	0.095665	31.28354	47.85613	0.6513
At most 3	0.034311	13.28406	29.79707	0.8782
At most 4	0.024933	7.034516	15.49471	0.5736
At most 5	0.013951	2.514875	3.841466	0.1128

Trace test indicates 1 cointegrating eqn(s) at the 0.05 level
* denotes rejection of the hypothesis at the 0.05 level
**MacKinnon-Haug-Michelis (1999) p-values

Unrestricted Cointegration Rank Test (Maximum Eigenvalue)

Hypothesized No. of CE(s)	Eigenvalue	Max-Eigen Statistic	0.05 Critical Value	Prob.**
None *	0.261502	54.26160	40.07757	0.0007
At most 1	0.157388	30.65350	33.87687	0.1156
At most 2	0.095665	17.99948	27.58434	0.4952
At most 3	0.034311	6.249549	21.13162	0.9774
At most 4	0.024933	4.519641	14.26460	0.8008
At most 5	0.013951	2.514875	3.841466	0.1128

Max-eigenvalue test indicates 1 cointegrating eqn(s) at the 0.05 level
* denotes rejection of the hypothesis at the 0.05 level
**MacKinnon-Haug-Michelis (1999) p-values

Table A12: Johansen's Cointegration Test (U.K.)

Unrestricted Cointegration Rank Test (Trace)

Hypothesized No. of CE(s)	Eigenvalue	Trace Statistic	0.05 Critical Value	Prob.**
None *	0.288398	156.5227	95.75366	0.0000
At most 1 *	0.261763	95.62031	69.81889	0.0001
At most 2	0.095609	41.29559	47.85613	0.1794
At most 3	0.089167	23.30732	29.79707	0.2313
At most 4	0.035447	6.589576	15.49471	0.6258
At most 5	0.000723	0.129439	3.841466	0.7190

Trace test indicates 2 cointegrating eqn(s) at the 0.05 level
* denotes rejection of the hypothesis at the 0.05 level
**MacKinnon-Haug-Michelis (1999) p-values

Unrestricted Cointegration Rank Test (Maximum Eigenvalue)

Hypothesized No. of CE(s)	Eigenvalue	Max-Eigen Statistic	0.05 Critical Value	Prob.**
None *	0.288398	60.90236	40.07757	0.0001
At most 1 *	0.261763	54.32472	33.87687	0.0001
At most 2	0.095609	17.98827	27.58434	0.4961
At most 3	0.089167	16.71774	21.13162	0.1857
At most 4	0.035447	6.460138	14.26460	0.5549
At most 5	0.000723	0.129439	3.841466	0.7190

Max-eigenvalue test indicates 2 cointegrating eqn(s) at the 0.05 level
* denotes rejection of the hypothesis at the 0.05 level
**MacKinnon-Haug-Michelis (1999) p-values

Table A13: Johansen's Cointegration Test (Japan)

Unrestricted Cointegration Rank Test (Trace)

Hypothesized No. of CE(s)	Eigenvalue	Trace Statistic	0.05 Critical Value	Prob.**
None *	0.256368	98.65954	79.34145	0.0009
At most 1	0.109911	45.34190	55.24578	0.2758
At most 2	0.088885	24.38386	35.01090	0.4207
At most 3	0.035967	7.628340	18.39771	0.7206
At most 4	0.005734	1.035024	3.841466	0.3090

Trace test indicates 1 cointegrating eqn(s) at the 0.05 level
* denotes rejection of the hypothesis at the 0.05 level
**MacKinnon-Haug-Michelis (1999) p-values

Unrestricted Cointegration Rank Test (Maximum Eigenvalue)

Hypothesized No. of CE(s)	Eigenvalue	Max-Eigen Statistic	0.05 Critical Value	Prob.**
None *	0.256368	53.31764	37.16359	0.0003
At most 1	0.109911	20.95804	30.81507	0.4752
At most 2	0.088885	16.75552	24.25202	0.3546
At most 3	0.035967	6.593316	17.14769	0.7566
At most 4	0.005734	1.035024	3.841466	0.3090

Max-eigenvalue test indicates 1 cointegrating eqn(s) at the 0.05 level
* denotes rejection of the hypothesis at the 0.05 level
**MacKinnon-Haug-Michelis (1999) p-values

Table A14: Johansen's Cointegration Test (Canada)

Unrestricted Cointegration Rank Test (Trace)

Hypothesized No. of CE(s)	Eigenvalue	Trace Statistic	0.05 Critical Value	Prob.**
None *	0.250217	134.8758	95.75366	0.0000
At most 1 *	0.202848	81.31301	69.81889	0.0046
At most 2	0.103499	39.14490	47.85613	0.2545
At most 3	0.080128	18.82341	29.79707	0.5057
At most 4	0.017138	3.288560	15.49471	0.9523
At most 5	0.000394	0.073234	3.841466	0.7867

Trace test indicates 2 cointegrating eqn(s) at the 0.05 level
* denotes rejection of the hypothesis at the 0.05 level
**MacKinnon-Haug-Michelis (1999) p-values

Unrestricted Cointegration Rank Test (Maximum Eigenvalue)

Hypothesized No. of CE(s)	Eigenvalue	Max-Eigen Statistic	0.05 Critical Value	Prob.**
None *	0.250217	53.56274	40.07757	0.0009
At most 1 *	0.202848	42.16811	33.87687	0.0041
At most 2	0.103499	20.32149	27.58434	0.3193
At most 3	0.080128	15.53485	21.13162	0.2532
At most 4	0.017138	3.215326	14.26460	0.9314
At most 5	0.000394	0.073234	3.841466	0.7867

Max-eigenvalue test indicates 2 cointegrating eqn(s) at the 0.05 level
* denotes rejection of the hypothesis at the 0.05 level
**MacKinnon-Haug-Michelis (1999) p-values

Table A15: Johansen's Cointegration Test (U.S.)

Appendix for Chapter 3: Model

We extend Bruno et al. (2018)'s model to analyze production chains with offshoring and its impact on emissions. First, we assume that the representative firm chooses whether to offshore tasks to another country or not. We assume that total production chain stages are $\bar{n}$. Total output (value-added) is assumed to be

$$\left(\sum_{i=1}^{\bar{n}} x_i\right)^{\alpha} \quad (0 < \alpha < 1), \tag{A.1}$$

where $x_i = 1 + b$ if a firm offshore the ith stage task. Otherwise, $x_i = 1$. Thus, if a firm offshores s stages to the foreign country, total revenue generated per worker is

$$y(\bar{n}; s) = (\bar{n} + bs)^{\alpha}. \tag{A.2}$$

Total emissions by the representative firm depends on the amount of output, which is directly affected by the extent of offshoring:

$$E = \left(\sum_{i=1}^{\bar{n}} e_i\right)^{\alpha} \quad (0 < \alpha < 1) \tag{A.3}$$

where $e_i = 1 - d$ if a firm offshore the ith stage task. Otherwise, $x_i = 1$. d refers to the damage on the environmental quality due to the production of outputs.[5] As a result, if the representative firm decides to offshore s stages to the countries with the absolute advantage, total emissions generated in the home country is

$$E(\bar{n}; s) = (\bar{n} - ds)^{\alpha}. \tag{A.4}$$

We assume that the representative firm faces an emission tax t per unit of emissions. This implies that the firm has additional incentives to offshore in order to minimize tax penalties.

Offshoring takes one unit of production chain stage as a transportation cost. For example, if a firm decides to offshore s tasks among $\bar{n}$ total stages to a foreign country, the total production process becomes $\bar{n} + s$. Total working capital for the whole process with s offshoring is

$$\begin{aligned} K &= \frac{1}{2}(\bar{n} + s)(\bar{n} + s + 1)\,\omega \times \frac{L}{(\bar{n} + s)} \\ &= \frac{\bar{n} + s + 1}{2}\omega L \end{aligned} \tag{A.5}$$

where L is the world labor force. Thus, the profit of the representative firm is

$$\Pi = (\bar{n} + bs)^{\alpha} zL - \omega zL - rzK - tE(s), \tag{A.6}$$

[5] We assume that emissions are still generated in the home country even after offshoring activities due to the transportation of goods to the foreign country.

where z is the proportion of the firm relative to the world production. From profit maximization problem and zero-profit condition of the representative firm, we are able to derive the following implicit function:

$$\begin{aligned}\Phi = &\left[\frac{2\alpha c}{r} + (1+\bar{n})\,\alpha c\right]\left[zL\,(\bar{n}+cs)^{\alpha-1} + t\,(\bar{n}-cs)^{\alpha-1}\right] \\ &+ s\alpha c\left[zL\,(\bar{n}+cs)^{\alpha-1} + t\,(\bar{n}-cs)^{\alpha-1}\right] - zL\,(\bar{n}+cs)^{\alpha} + t\,(\bar{n}-cs)^{\alpha}\,, \end{aligned} \tag{A.7}$$

where c is assumed to be equal to b and d for analytical simplicity.
Equation (A.7) enables us to conduct a comparative static. Since the main theme of this paper is the explore the impact of financing costs—which is affected by the monetary policy by the government—on the environmental quality, we aim to show the sign of $\frac{\partial s}{\partial r}$. From the implicit function theorem, we need to compute

$$\frac{\partial s}{\partial r} = -\frac{\frac{\partial \Phi}{\partial r}}{\frac{\partial \Phi}{\partial s}}. \tag{A.8}$$

It is clear that

$$\frac{\partial \Phi}{\partial r} < 0. \tag{A.9}$$

However, the sign of the denominator of Equation (A.8) is ambiguous depends upon the relative significance of the parameters. If the environmental regulations are sufficiently loose such that $zL \gg t$, we have

$$\frac{\partial \Phi}{\partial s} = (-) + ({}^{+}\!/\!_{-}) + (-) + (-), \tag{A.10}$$

which implies that the sign is likely to be negative. However, if the home country imposes a stringent environmental regulations where $zL \ll t$, then we have

$$\frac{\partial \Phi}{\partial s} = (+) + (+) + (-) + (-), \tag{A.11}$$

which indicates that the sign is likely to be positive. These results lead us to derive a conclusion that there are chances that greater credit cost induces more offshoring if environmental regulation is sufficiently large, and vice versa, otherwise. Considering the fact that

$$\frac{\partial E}{\partial r} = \underbrace{\frac{\partial E}{\partial s}}_{(-)}\,\underbrace{\frac{\partial s}{\partial r}}_{(+/-)}\,, \tag{A.12}$$

we can infer that the monetary policy which induces stronger dollars may still cause North countries (i.e., which are likely to impose strict regulations) to offshore more and thus improve the air quality. However, this is not the case for the South countries. Due to their loose environmental regulations, the representative firm has less incentives to offshore in response to stronger dollars, which in turn raise the stock of total emissions.

Derivation of Equation (A.7)

From profit maximization problem, the extent of offshoring $s$is determined. The first-order condition for s gives

$$
\begin{aligned}
\frac{\partial \Pi}{\partial s} &= \alpha\left(\bar{n}+bs\right)^{\alpha-1} bzL - \frac{\omega zLr}{2} - tE'(s) = 0, \\
&\Rightarrow \left(\bar{n}+bs\right)^{\alpha-1} = \left[\frac{\omega zLr}{2} + tE'(s)\right]/\alpha bzL, \\
&\Rightarrow \left(\bar{n}+bs\right)^{\alpha-1} = \frac{\omega r}{2\alpha b} + \frac{tE'(s)}{\alpha bzL} \\
&\Rightarrow \left(\bar{n}+bs\right) = \left(\frac{\omega r}{2\alpha b} + \frac{tE'(s)}{\alpha bzL}\right)^{\frac{1}{\alpha-1}} \\
&\Rightarrow \left(\bar{n}+bs\right) = \left(\frac{\omega rzL}{2\alpha bzL} + \frac{2tE'(s)}{2\alpha bzL}\right)^{\frac{1}{\alpha-1}} = \left(\frac{\omega rzL + 2tE'(s)}{2\alpha bzL}\right)^{\frac{1}{\alpha-1}} = \left(\frac{2\alpha bzL}{\omega rzL + 2tE'(s)}\right)^{\frac{1}{1-\alpha}} \\
&\Rightarrow \left(\bar{n}+bs\right)^{1-\alpha} = \frac{2\alpha bzL}{\omega rzL + 2tE'(s)} \\
&\Rightarrow \omega rzL + 2tE'(s) = \frac{2\alpha bzL}{\left(\bar{n}+bs\right)^{1-\alpha}} \\
&\Rightarrow \omega = \frac{1}{rzL}\left[\frac{2\alpha bzL}{\left(\bar{n}+bs\right)^{1-\alpha}} - 2tE'(s)\right] \\
&\Rightarrow \omega = \frac{1}{rzL}\left[\frac{2\alpha bzL}{\left(\bar{n}+bs\right)^{1-\alpha}} + 2dt\alpha\left(\bar{n}-ds\right)^{\alpha-1}\right] \qquad \text{(A.13)}
\end{aligned}
$$

The zero profit condition gives

$$
\begin{aligned}
\Pi &= \left(\bar{n}+bs\right)^{\alpha} zL - \omega zL - rzK - tE(s) = 0, \\
&\Rightarrow \left(\bar{n}+bs\right)^{\alpha} zL = \omega zL\left(1 + \frac{r\left(\bar{n}+s+1\right)}{2}\right) + tE(s) \\
&\Rightarrow \left(\bar{n}+bs\right)^{\alpha} = \omega\left(1 + \frac{r}{2}\left(1+\bar{n}\right) + \frac{r}{2}s\right) + \frac{tE(s)}{zL} \\
&\Rightarrow \left(\bar{n}+bs\right)^{\alpha} = \omega\left(1 + \frac{r}{2}\left(1+\bar{n}\right) + \frac{r}{2}s\right) + \frac{t\left(\bar{n}-ds\right)^{\alpha}}{zL} \qquad \text{(A.14)}
\end{aligned}
$$

Equations (A.13) and (A.14) give

$$
\begin{aligned}
\left(\bar{n}+bs\right)^{\alpha} &= \frac{1}{rzL}\left[\frac{2\alpha bzL}{\left(\bar{n}+bs\right)^{1-\alpha}} + 2dt\alpha\left(\bar{n}-ds\right)^{\alpha-1}\right]\left(1 + \frac{r\left(1+\bar{n}\right)}{2} + \frac{rs}{2}\right) + \frac{t\left(\bar{n}-ds\right)^{\alpha}}{zL} \\
&= \left[\frac{2\alpha b}{r\left(\bar{n}+bs\right)^{1-\alpha}} + \frac{2dt\alpha\left(\bar{n}-ds\right)^{\alpha-1}}{rzL}\right]\left(\frac{2+r\left(1+\bar{n}\right)}{2} + \frac{rs}{2}\right) + \frac{t\left(\bar{n}-ds\right)^{\alpha}}{zL}
\end{aligned}
$$

$$
\begin{aligned}
1 &= \frac{1}{\left(\bar{n}+bs\right)^{\alpha}}\left\{\left[\frac{2\alpha b}{r\left(\bar{n}+bs\right)^{1-\alpha}} + \frac{2dt\alpha\left(\bar{n}-ds\right)^{\alpha-1}}{rzL}\right]\left(\frac{2+r\left(1+\bar{n}\right)}{2} + \frac{rs}{2}\right) + \frac{t\left(\bar{n}-ds\right)^{\alpha}}{zL}\right\} \\
&= \left[\frac{2\alpha b}{r\left(\bar{n}+bs\right)} + \frac{2dt\alpha\left(\bar{n}-ds\right)^{\alpha-1}}{rzL\left(\bar{n}+bs\right)^{\alpha}}\right]\left(\frac{2+r\left(1+\bar{n}\right)}{2} + \frac{rs}{2}\right) + \frac{1}{\left(\bar{n}+bs\right)^{\alpha}}\frac{t\left(\bar{n}-ds\right)^{\alpha}}{zL}
\end{aligned}
$$

$$
\begin{aligned}
1-\frac{1}{(\bar{n}+bs)^{\alpha}}\frac{t(\bar{n}-ds)^{\alpha}}{zL} &= \left[\frac{2\alpha b}{r(\bar{n}+bs)}+\frac{2dt\alpha(\bar{n}-ds)^{\alpha-1}}{rzL(\bar{n}+bs)^{\alpha}}\right]\left(\frac{2+r(1+\bar{n})}{2}+\frac{rs}{2}\right)\\
\frac{zL(\bar{n}+bs)^{\alpha}-t(\bar{n}-ds)^{\alpha}}{zL(\bar{n}+bs)^{\alpha}} &= \left[\frac{2\alpha b}{r(\bar{n}+bs)}+\frac{2dt\alpha(\bar{n}-ds)^{\alpha-1}}{rzL(\bar{n}+bs)^{\alpha}}\right]\left(\frac{2+r(1+\bar{n})}{2}+\frac{rs}{2}\right)\\
zL(\bar{n}+bs)^{\alpha}-t(\bar{n}-ds)^{\alpha} &= zL(\bar{n}+bs)^{\alpha}\left[\frac{2\alpha b}{r(\bar{n}+bs)}+\frac{2dt\alpha(\bar{n}-ds)^{\alpha-1}}{rzL(\bar{n}+bs)^{\alpha}}\right]\left(\frac{2+r(1+\bar{n})}{2}+\frac{rs}{2}\right)\\
&= \left[\frac{2\alpha bzL(\bar{n}+bs)^{\alpha-1}}{r}+\frac{2dt\alpha(\bar{n}-ds)^{\alpha-1}}{r}\right]\left(\frac{2+r(1+\bar{n})}{2}+\frac{rs}{2}\right)
\end{aligned}
$$

For simplicity, we assume that $b=d=c$. Then,

$$
\begin{aligned}
zL(\bar{n}+cs)^{\alpha}-t(\bar{n}-cs)^{\alpha} &= \left[\frac{2\alpha czL(\bar{n}+cs)^{\alpha-1}}{r}+\frac{2ct\alpha(\bar{n}-cs)^{\alpha-1}}{r}\right]\left(\frac{2+r(1+\bar{n})}{2}+\frac{rs}{2}\right)\\
&= 2\alpha c\left[\frac{zL(\bar{n}+cs)^{\alpha-1}}{r}+\frac{t(\bar{n}-cs)^{\alpha-1}}{r}\right]\left(\frac{2+r(1+\bar{n})}{2}+\frac{rs}{2}\right)
\end{aligned}
$$

$$
\begin{aligned}
\mathit{\Phi} &\equiv 2\alpha c\left[\frac{zL(\bar{n}+cs)^{\alpha-1}}{r}+\frac{t(\bar{n}-cs)^{\alpha-1}}{r}\right]\left(\frac{2+r(1+\bar{n})}{2}+\frac{rs}{2}\right)-zL(\bar{n}+cs)^{\alpha}+t(\bar{n}-cs)^{\alpha}\\
&= 2\alpha c\left[\frac{zL(\bar{n}+cs)^{\alpha-1}}{r}+\frac{t(\bar{n}-cs)^{\alpha-1}}{r}\right]\left(1+\frac{r(1+\bar{n})}{2}+\frac{rs}{2}\right)-zL(\bar{n}+cs)^{\alpha}+t(\bar{n}-cs)^{\alpha}\\
&= 2\alpha c\left[\frac{zL(\bar{n}+cs)^{\alpha-1}+t(\bar{n}-cs)^{\alpha-1}}{r}\right]\left(1+\frac{r(1+\bar{n})}{2}+\frac{rs}{2}\right)-zL(\bar{n}+cs)^{\alpha}+t(\bar{n}-cs)^{\alpha}\\
&= 2\alpha c\left[\frac{zL(\bar{n}+cs)^{\alpha-1}+t(\bar{n}-cs)^{\alpha-1}}{r}\right]+\left[zL(\bar{n}+cs)^{\alpha-1}+t(\bar{n}-cs)^{\alpha-1}\right](1+\bar{n})\alpha c\\
&+\left[zL(\bar{n}+cs)^{\alpha-1}+t(\bar{n}-cs)^{\alpha-1}\right]s\alpha c-zL(\bar{n}+cs)^{\alpha}+t(\bar{n}-cs)^{\alpha}\\
\mathit{\Phi} &= \left[\frac{2\alpha c}{r}+(1+\bar{n})\alpha c\right]\left[zL(\bar{n}+cs)^{\alpha-1}+t(\bar{n}-cs)^{\alpha-1}\right]+s\alpha c\left[zL(\bar{n}+cs)^{\alpha-1}+t(\bar{n}-cs)^{\alpha-1}\right]\\
&-zL(\bar{n}+cs)^{\alpha}+t(\bar{n}-cs)^{\alpha}\,. \qquad \text{(A.15)}
\end{aligned}
$$

Detailed Comparative Static of Equation (A.7)

The impact of credit costs on the extent of offshoring, $\frac{\partial s}{\partial r}$, is determined by the implicit function theorem:

$$\frac{\partial s}{\partial r} = -\frac{\frac{\partial \Phi}{\partial r}}{\frac{\partial \Phi}{\partial s}}.$$

It is clear that

$$\frac{\partial \Phi}{\partial r} = -2\alpha c r^{-2}\left[zL\left(\bar{n}+cs\right)^{\alpha-1} + t\left(\bar{n}-cs\right)^{\alpha-1}\right] < 0.$$

This sign of the denominator, however, is ambiguous:

$$\frac{\partial \Phi}{\partial s} \lesseqgtr 0,$$

since the sign of the first term of the implicit function A.15 is determined by the following derivative,

$$\begin{aligned}\frac{\partial\left[zL\left(\bar{n}+cs\right)^{\alpha-1} + t\left(\bar{n}-cs\right)^{\alpha-1}\right]}{\partial s} &= (\alpha-1)zL\left(\bar{n}+cs\right)^{\alpha-2}c - c(\alpha-1)t\left(\bar{n}-cs\right)^{\alpha-2} \\ &= \underbrace{(\alpha-1)c}_{(-)}\underbrace{\left[zL\left(\bar{n}+cs\right)^{\alpha-2} - t\left(\bar{n}-cs\right)^{\alpha-2}\right]}_{(+/-)} \lesseqgtr 0. \qquad \text{(A.16)}\end{aligned}$$

The sign of Equation (A.16) is determined by the relative significance of zL and t. Also, the sign of the second term of the implicit function is also affected by the relative significance of zL and t:

$$\begin{aligned}\frac{\partial s\alpha c\left[zL\left(\bar{n}+cs\right)^{\alpha-1} + t\left(\bar{n}-cs\right)^{\alpha-1}\right]}{\partial s} &= \underbrace{\alpha c\left[zL\left(\bar{n}+cs\right)^{\alpha-1} + t\left(\bar{n}-cs\right)^{\alpha-1}\right]}_{(+)} \\ &+ \underbrace{s\alpha c(\alpha-1)c\left[zL\left(\bar{n}+cs\right)^{\alpha-2} - t\left(\bar{n}-cs\right)^{\alpha-2}\right]}_{(+/-)}. \qquad \text{(A.17)}\end{aligned}$$

If $zL \ll t$, the first two terms of the implicit function become positive, while if $zL \gg t$, only the sign of the second term becomes ambiguous.