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Independence + Accountability: Why the Fed Is a Well-Designed Central Bank

Christopher J. Waller

In 1913, Congress purposefully created the Federal Reserve as an independent central bank, which created a fundamental tension: how to ensure the Fed remains accountable to the electorate without losing its independence. Over the years, there have been changes in the Fed's structure to improve its independence, credibility, accountability, and transparency. These changes have led to a better institutional design that makes U.S. policy credible and based on sound economic reasoning, as opposed to politics. In times of financial and economic crisis, there is an understandable tendency to reexamine the structure of the Federal Reserve System. A central bank's independence, however, is the key tool to ensure a government will not misuse monetary policy for short-term political reasons. (JEL E52, E58)

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The Federal Reserve has taken unprecedented actions in the financial markets since the advent of the financial crisis. Noteworthy examples include lending more than $1.5 trillion to financial institutions and buying $1.25 trillion of mortgage-backed securities to stabilize the economy. The large scale of these interventions has brought intense public scrutiny of the Federal Reserve's powers and institutional structure. In particular, many have questioned why the Fed has the freedom to engage in such actions without the explicit consent from Congress or the president. This freedom from political interference is commonly referred to as "central bank independence."

The focus of this article is to review why Congress made the Federal Reserve independent when it created it in 1913. The article also addresses the fundamental tension that comes with an independent central bank: how to ensure that these policymakers are accountable to the electorate without losing their independence. The key point to remember is that giving the central bank independence is the best method for governments to tie their own hands and prevent them from misusing monetary policy for short-term political reasons.

THE POWER OF MONEY

Money is obviously a vital part of an economy because it allows trade to occur more efficiently. Governments have a great power that no one else in the economy has—the ability to print money. Thus, the government can acquire more goods by printing more money, a process known as seigniorage. This power, however, brings with it a dangerous temptation. Imagine that you had this power; just think of what you could do with it! You could live a great life, feed the hungry, and house the homeless. And all of this could be achieved simply by printing more money. This sounds wonderful. How can it be dangerous?

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If the government prints too much money, people who sell things for money raise their prices. (These prices can apply to goods, services, and labor.) This lowers the purchasing power and value of the money being printed. In fact, if the government prints too much money, the money becomes worthless. We have seen many governments give in to this temptation, and the result is a hyperinflation. Hyperinflations were observed in the 20th century in Germany (twice), Hungary, Ecuador, Bolivia, and Peru, with Zimbabwe as the most recent casualty. Such episodes of high inflation can greatly impair the functioning of the economy or collapse it altogether. Thus, having the power to print money brings with it great responsibility to respect that power.

It is important to remember that the temptation to print money is not restricted to less-developed countries. In fact, the United States has suffered from high inflation several times. In pre-revolutionary days, many colonies had the right to print money and fell prey to their own excesses. The Continental Congress did the same during the Revolutionary War. In 1775, it gave the colonies the authority to issue Continental dollars to finance the war. Overissuance and counterfeiting by the British led to such dramatic increases in paper currency that by 1779, the value of a Continental dollar was 1/25th of its original value (giving rise to the phrase "not worth a continental"). During the Civil War, the Confederate government also succumbed to the temptation of printing money to buy goods. From 1861 to 1864, the stock of Confederate dollars increased 10-fold, and prices increased the same. Financing government spending via the printing press also occurred in the 20th century. Shortly after the founding of the Federal Reserve, the U.S. Treasury adopted policies that induced the Fed to monetize government debt.[1] This led to a spike in U.S. inflation following World War I. These examples show that the U.S. government has a history of resorting to the printing press to pay for government expenditures.

Most governments have taken steps to discipline themselves and impose restraints on their ability to print money to pay for goods. A time-honored method of restraint was to tie the value of the currency to a commodity such as gold. Because the government did not control gold production, the amount of money it could print was limited by its holdings of gold. Although this restrained the government's ability to create seigniorage, it also unfortunately tied its hands during periods of high demand for currency, such as financial crises (a time in which people wanted to hold the government's currency rather than other assets) or during planting season (a time in which farmers needed cash to pay for seed, etc.). Other problems also occurred: New gold discoveries, such as during the California gold rush, led to an inflow of gold and new currency issue, which caused inflation. Conversely, if the economy grew faster than the supply of gold, then prices of goods and services would fall, leading to deflation. Finally, it is very costly to mine gold simply to hold it in storage to back up pieces of paper money. For these reasons and others, governments began to realize that using a gold standard to control the nation's money supply was too restrictive and costly.

As a result, governments slowly moved to a fiat currency system, one in which the money was not backed by a commodity but rather by the "full faith and credit" of the government. Under such a system, the government promises its citizens that it will discipline itself and not resort to seigniorage to finance government spending. In short, citizens have to trust that the government will do the right thing. But trust can be abused; therefore, the citizenry demanded institutional arrangements that backed up the government's pledge.

That is why most governments took steps to tie their own hands and make themselves credible stewards of their nation's economic interests. It became very clear that if elected government officials had direct control of the money supply, then they could cut taxes and print money to pay for goods to win votes. Consequently, promises by elected officials would not be seen as credible. To achieve credibility and avoid this abuse of

[1] Monetizing debt means the government borrows money to buy goods and then repays its debt by printing more money. This is equivalent to simply printing money in the first place to buy goods.

public power for private gain, the control of the money supply had to be delegated to a nonelected group of individuals. These officials were to run the institution responsible for monetary policy, known as the "central bank."

It has always been important that central bankers be independent of the political process to ensure that they cannot be manipulated by elected officials. However, having such great power means that central bankers have to be accountable to the electorate in some fashion, and accountability requires the central bank to behave in a transparent manner. Thus a well-designed central bank needs to be (i) credible, (ii) independent, (iii) accountable, and (iv) transparent.

CENTRAL BANK INDEPENDENCE AND INFLATION

A key macroeconomic axiom is that sustained high growth rates of a nation's money stock in excess of its production of goods and services eventually produce high and rising inflation rates. This axiom was nicely phrased by Milton Friedman when he said that "inflation is always and everywhere a monetary phenomenon." Economic history is littered with countries that ran afoul of this axiom. A recent example is Zimbabwe, which saw its annual inflation rate rise from 24,411 percent in 2007 to an estimated 89.7 sextillion percent in mid-November 2008.[2] That's 89,700,000,000,000,000,000,000 percent.

The willingness of governments to force their central banks to print excessive amounts of money, or put in place policies that lead to higher inflation rates over time, has been termed the "inflation bias" of discretionary monetary policymaking. (See Walsh, 2008.) To minimize this bias, many governments have decided to give their central bank legal independence (CBI). But do countries with independent central banks also have lower inflation? To answer this question properly, it's necessary to calculate country-specific measures of central bank independence. Many economists have constructed measures of CBI from a variety of legal indicators, many of which are discussed in this article. In a now famous article that was published in 1993, Alesina and Summers (1993) found that developed countries with high levels of central bank independence also experienced lower average levels of inflation for the period 1955-88. Figure 1 is derived from a figure in their paper, which clearly shows this negative relationship.

More recently, as the top chart in Figure 2 shows, global inflation has slowed sharply since the mid-1990s. However, as the bottom two charts indicate, the rapid descent in global inflation was due primarily to conditions in emerging market and developing countries. In the developed countries, the slowing occurred much earlier, in the early 1980s. There were many reasons for the global decline in inflation since the late 1980s, including stronger commitments to price stability (better monetary policies), higher rates of productivity growth, and the forces of globalization that increased competition and enhanced the flexibility of labor and product markets. (See Rogoff, 2003.) As suggested by Alesina and Summers, increased central bank independence appears to be another key reason for the decline in inflation worldwide. As shown in Table 1, there was a marked increase in central bank independence between the period 1980-89 and 2003.

Although this trend was apparent among developed countries, it was especially apparent among emerging market and developing countries.[3] Indeed, many of the reforms that enhanced central bank independence occurred during the 1990s and were in response to high rates of inflation. (See Cukierman, 2008.) The movement toward greater central bank independence undoubtedly helps to explain the sharp slowing in inflation in many countries.

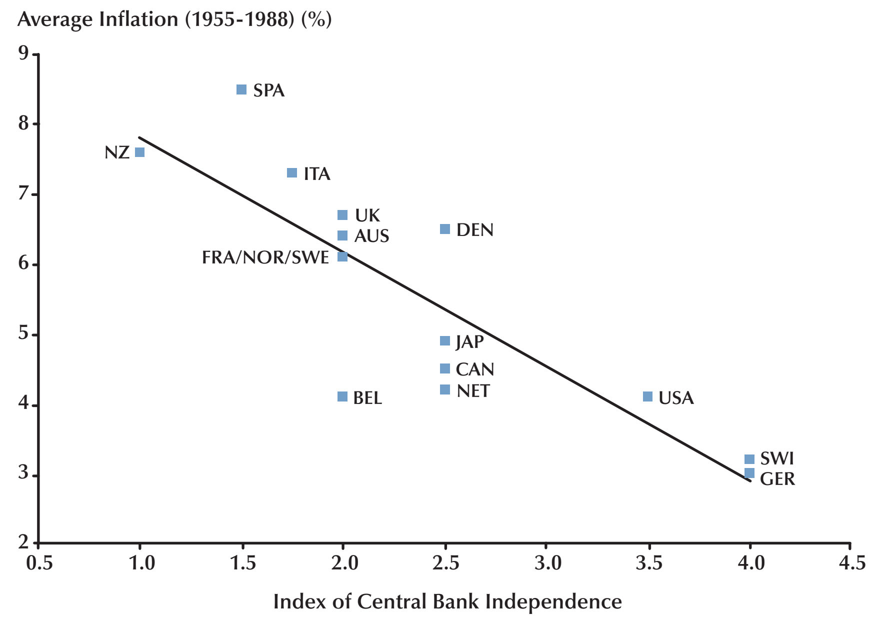
There was also an increase in CBI in advanced countries. However, the overall movement from weak and moderate independence to strong independence arose mostly from those countries that joined the European Union and thus became members of the European Central Bank (ECB). Because of the Maastricht Treaty, the ECB is

[2] See Hanke and Kwok (2009).

[3] The data are published in Crowe and Meade (2007).

Figure 1

Central Bank Independence versus Average Inflation



NOTE: Derived from Alesina and Summers (1993).

Table 1

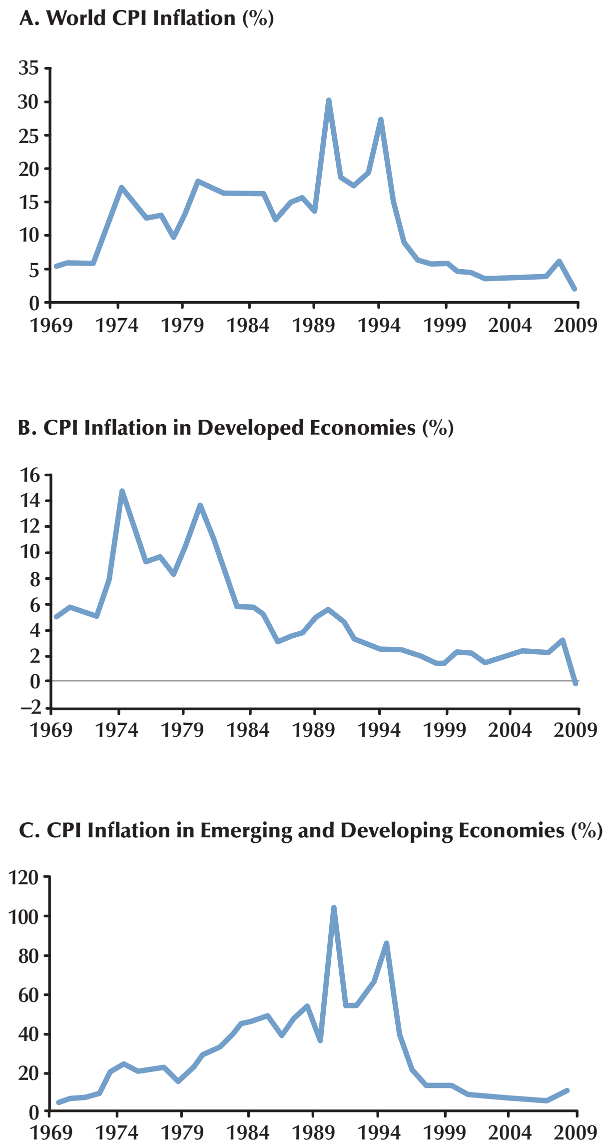
Measures and Frequency Distribution of Central Bank Independence

	Developed Economies			Emerging & Developing Economies		
	1980-89	**2003**	**Net Change**	**1980-89**	**2003**	**Net Change**
Weak independence	13	8	–5	32	6	–26
Moderate independence	8	5	–3	19	49	30
Strong independence	0	13	13	0	15	15

NOTE: Crowe and Meade (2007) measure central bank independence on a numerical scale from 0 (no independence) to 1 (complete independence). For this table, weak CBI is defined to include those banks with a measure from 0 to less than 0.4; moderate CBI is defined as those banks from 0.4 to 0.8; strong CBI is for banks with a measure of 0.8 or above. The Federal Reserve's ranking on this scale is 0.47, and the ECB's ranking is 0.83.

Figure 2

Central Bank Independence versus Average Inflation



SOURCE: International Monetary Fund.

deemed to be strongly independent. Interestingly, while the trend over the past 20 years or so is toward increasing CBI, the Federal Reserve has not become more independent, according to the measure shown in the table. Still, the U.S. inflation rate has slowed markedly since the 1970s and 1980s. This suggests that CBI may be necessary but not sufficient to produce good inflation performance over time—a result that seems to hold for other developed countries as well. However, central bank independence seems to have been much more important for helping to explain the sharp decline in inflation rates since the 1980s for emerging market and developing economies.

A SERIES OF CHECKS AND BALANCES

The tricky issue is that accountability means being subject to some political oversight, which weakens the perception that the central bank is independent. So, there is an inherent tension between having independence to conduct policy and being accountable to the electorate. Furthermore, if central bankers are not elected, then they must be chosen in another way. But by whom?

In the United States, there has long been a tension between the states and the federal government. States were leery of giving too much power to the federal government out of fear that this power would be abused. Yet, the federal government was the body charged with the welfare of the entire nation. In response to this conflict between the states and the federal government, a series of checks and balances was implemented to ensure that policy was conducted in a way that protected both interests. So, it is not surprising that similar checks and balances would come into play when deciding who selects the nonelected officials to run monetary policy and to whom they would be accountable. Thus, while the Federal Reserve was created to conduct monetary policy, it was given a complicated system of checks and balances to deal with conflicts between the states and the federal government, as well as between the legislative and executive branches of the federal government.

What are these checks and balances? First, rather than have a single central bank, the founders created a system of central banks. This system includes the Board of Governors in Washington, D.C., and 12 regional Reserve Banks. This arrangement avoided the problem of having strong federal government control of the central bank. The idea behind the regional banks is that the further these policymakers are from the day-to-day political process, the more likely that monetary policy decisions would be made on economic grounds rather than political considerations. Furthermore, the policymakers would be less susceptible to pressures to create seigniorage. The opposite concern is that the regional Banks would focus too much on their own regions (or, in Fed parlance, Federal Reserve Districts). Therefore, the Board of Governors (seven members) was created to ensure that the entire nation's welfare was considered. Thus, policy was to be set by the 12 presidents of the regional Banks (those who served as direct contacts with the states) and the seven members of the Board of Governors (those who were intended to have more of a national view).

Second, who would choose these 19 policymakers? One concern of the founders was that if all of the central bankers were political appointees of the president or Congress, then the Fed would not have the independence it needed to conduct policy in an appropriate manner. It therefore was decided that the presidents of the regional banks would not be political appointees but would be chosen by the citizenry of their Districts in a non-electoral manner. This ensured that the presidents would be independent of the political process and less likely to engage in seigniorage creation. One method of choosing regional presidents in a non-electoral manner was to create a local board of directors for each of the 12 regional Reserve Banks. Each board, in turn, would select its regional Bank president. To achieve a broad perspective on the economic well-being of each District, the board was to be composed of individuals from a wide range of sectors. This ensured that the regional Bank presidents would be chosen based on their professional qualifications as opposed to their political connections or sectoral ties.

On the other hand, because 12 of the 19 policymakers were not political appointees, there was concern that there was not enough accountability to the electorate. Thus, it was decided that the seven members of the Board of Governors should be political appointees. The president would have the power to nominate the governors, and the Senate would have the power to confirm them. Consequently, this procedure for selecting the 19 central bankers of the Federal Reserve System provided for both independence and accountability.

Third, a common method for politicians to entice government agencies to carry out specific political agendas is to threaten to cut the agencies' budgets. Thus, no matter how far the presidents of the regional Banks were from Washington, D.C., or how they were chosen, if the Federal Reserve did not have budget autonomy, then Congress could always threaten to cut its budget to get the Fed to carry out monetary policies that Congress desired. This power of the purse strings would undermine the Fed's independence and credibility to keep money creation low and stable. To counteract this possibility, Congress gave the Federal Reserve budget autonomy when it created the Fed in 1913. The Fed was given the power to earn its own income and spend it without government interference.[4] However, recognizing that the Fed was creating seigniorage for the nation as a whole, Congress directed the Fed to return any excess income to the federal government. To guarantee that excess income was returned, the Fed's income statement and balance sheet had to be transparent and auditable, not by Congress, but by an independent auditing agency to prevent political machinations. Again, checks and balances prevailed.

Fourth, to ensure the credibility of Fed promises to keep money creation under control, Congress created long terms of office for the Board of Governors (14 years) and staggered the governors' terms (one expires every two years). This effectively guaranteed that one president could not appoint all of the members of the Board

[4] It is interesting to note that, in effect, the members of Congress in 1913 ensured that Congress could not threaten the Fed with budget cuts in the future. Thus, an earlier generation of politicians implemented checks and balances on future generations of congressional representatives.

and therefore "stack" the Fed. Long terms also made the Board more independent of the political process because members did not have to worry about reappointment. Finally, long terms made the Board members more accountable: Policymakers who made promises today would likely still be in office in the future and could be brought to task for failing to live up to earlier promises. As a result, long terms gave current Board members an incentive to carry out promises.

Last, to prevent the Fed from making decisions that benefited a particular industry or region, Congress required the Fed to report on its actions. But to ensure that the Fed maintained its independence, Congress restrained itself from making frequent intrusions. The Fed was therefore required to report regularly to Congress; in return, Congress would not try to influence Fed decisions on a day-to-day or month-to-month basis. This reporting structure again gave the Fed independence, yet made it accountable and transparent to the electorate.

WILL THE FINANCIAL CRISIS FURTHER LIMIT THE FED'S INDEPENDENCE? SHOULD IT?

The recent recession and financial crisis were, in many respects, the worst since the 1930s.[5] In response, some economists and policymakers have begun to examine the Fed's policies prior to and during the financial crisis to see whether its goals, responsibilities, or institutional structure should be changed to help prevent another financial calamity.

The Federal Reserve Act of 1913 was designed to balance the competing interests of the public and private sectors. Some were afraid of excessive government intervention in private capital markets, while others were worried that the financial sector would have too much influence on the nation's economic well-being. In this spirit, the Act also sought to balance the interests of Wall Street (financial) and Main Street (business and agriculture). This system, by and large, has served the country well.

Fast forward to 2011. In response to the financial crisis and recession, some people argue that power should be further consolidated in Washington, D.C., to avoid another financial calamity. However, as St. Louis Fed president James Bullard and other Federal Reserve officials and private-sector economists have pointed out, moving the levers of monetary policy even closer to the hub of politics could lead to an erosion of the Fed's independence and, eventually, poor economic performance.[6]

Clearly, part of the desire to subject the Federal Reserve to greater political oversight is natural in a democracy—and may even be a healthy rebalancing to correct misplaced priorities or policies. Few would quibble with the argument that, in a democracy, central banks should be held accountable for their policies. Indeed, if the central bank puts in place policies that run counter to its stated goals, then that will damage the credibility of the central bank. And to a central bank, credibility is something that is valued highly. If a central bank's policies are not credible, then the bank will eventually lose the support of the nation's policymakers—and maybe its independence.

As part of the Fed's accountability to the public, senior Federal Reserve officials testify regularly before Congress. As Figure 3 shows, the number of congressional appearances by Federal Reserve officials has increased significantly over the past few years. This development is probably not too surprising, given the recent financial market turbulence. In addition, appearances by Federal Reserve officials tend to be higher during recessions, as in the early 1980s and the early 1990s. Although part of the increase in congressional appearances over time may reflect a general increase in the number of hearings, it is nonetheless clear that Congress actively scrutinizes the Fed's policies during times of tranquility as well as turmoil. The number of appearances over the

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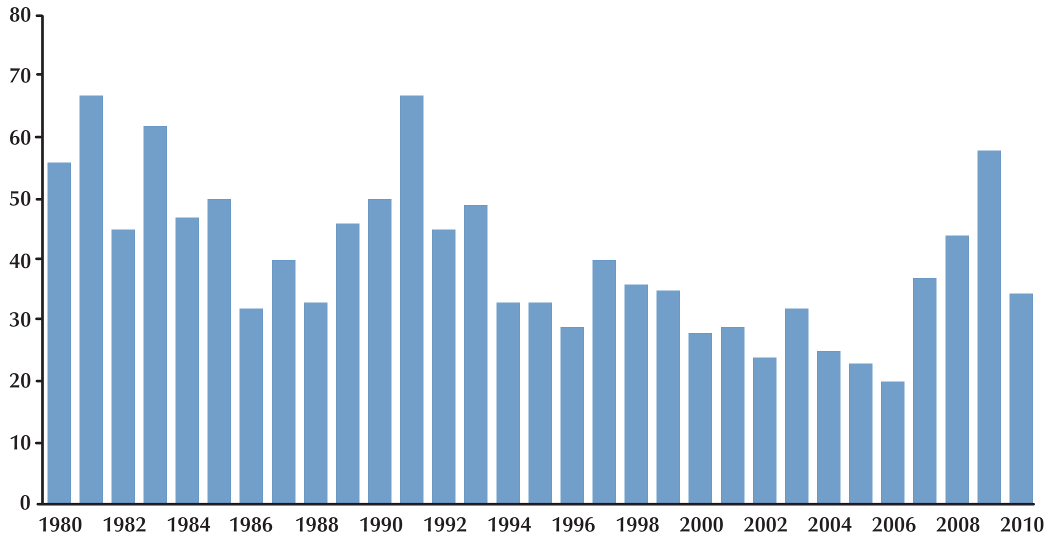
5 The causes and consequences of the financial crisis have been studied in depth. See the collection of articles and papers listed on the St. Louis Fed's financial crisis timeline at http://timeline.stlouisfed.org/index.cfm?p=articles.

6 See "The Fed at a Crossroads" (http://research.stlouisfed.org/econ/bullard/BullardWinterInstituteFinal.pdf) by James Bullard, president of the Federal Reserve Bank of St. Louis.

Figure 3

Congressional Appearances and Testimonies by Federal Reserve Officials, 1980-2010



SOURCE: Data compiled by the Federal Reserve Bank of St. Louis using the ProQuest Congressional database.

past four years is on pace to be the largest in about 20 years.

CONCLUSION

Over the years, there have been changes in the Fed's structure to improve its independence, credibility, accountability, and transparency. These changes have led to a better institutional design that makes U.S. policy credible and based on sound economic reasoning, as opposed to politics. In times of financial and economic crisis, there is a tendency to reexamine the structure of the Federal Reserve System. To the uninformed observer, the Fed's structure is in many ways mind-boggling. In particular, it seems counterintuitive that, in a democracy, the central bank should have independence from Congress. Yet, this independence is the result of Congress trying to avoid making monetary policy mistakes for political gain. Of course, accountability of public policymakers is a fundamental principle in a democracy. It is the tension between independence and accountability that led to the design of the Federal Reserve, and it has been an ever-present force in U.S. monetary policy for the past century.

In the end, the Federal Reserve System is a well-designed institution, created by Congress, that keeps the government from relying on the printing press to finance public spending. It is independent, credible, accountable, and transparent. It is a nearly 100-year-old success story that has served the nation well.

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A Foreign Exchange Intervention in an Era of Restraint

Christopher J. Neely

The Japanese yen appreciated strongly and rapidly against other major currencies in the wake of the massive March 11, 2011, Tohoku earthquake. High volatility and disorder in financial markets prompted the G-7 authorities to jointly intervene to weaken the yen. This episode resembled the two most recent G-7 coordinated interventions: the June 1998 effort to strengthen the yen and the September 2000 effort to strengthen the euro. Exchange rates reacted strongly and quickly to these three interventions, moving 3 to 4 percent in the desired direction within 30 minutes of the announcement and exhibiting lower volatility in the following days. G-7 authorities have used intervention very sparingly since 1995, yet the March 2011 policy action is a reminder that it can be used to calm markets and move the exchange rate in the desired direction. Intervention has become much less common but more successful. (JEL F31, E44, E58)

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An enormous, 9.0-magnitude earthquake rocked Japan on March 11, 2011, unleashing a tsunami that swamped the Japanese coast, killing more than 15,000 people and causing hundreds of billions of dollars of property damage.[1] This tremendous shock created great uncertainty in Japanese financial markets, raising concerns about future international trade and capital flows, increased expectations of government default, and reduced equity prices.

The Japanese yen (JPY) rapidly appreciated in the wake of the earthquake; from March 10 to March 17, 2011, its value rose by about 5 percent against the U.S. dollar (USD).[2] The financial press cited two factors contributing to the yen's rise: (i) expectations that Japanese insurance companies would need to liquidate and repatriate reserves held as foreign assets and (ii) the closing of "carry trade" positions in which investors borrowed in yen to lend abroad. In the days after the earthquake, the foreign currency and equity markets became extremely volatile.

In response to these volatile market conditions, the G-7 finance ministers and central bank governors announced late on Thursday, March 17, that they would jointly intervene the next day to reduce the value of the yen. The G-7 authorities cited concerns about "excess volatility and disorderly movements" in their intervention press release. In response, the yen depreciated by 3 to 4 percent—depending on the exchange rate—within hours.

This unusual intervention received limited media coverage. Articles in the financial press necessarily provided superficial coverage in the available space (see Pett, 2011; Vieira, 2011b;

1 See Hosaka (2011) and National Police Agency of Japan (2011).

2 The JPY/USD exchange rate fell from 82.98 JPY/USD at noon U.S. eastern time on March 10 to 78.74 JPY/USD at the same time on March 17, 2011.

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McCormick, 2011). Sack and McNeil (2011) briefly, but very usefully, summarized the facts from the policymaker's point of view. The rarity of such episodes since 1995 and the limited press coverage will leave many readers unfamiliar with foreign exchange intervention as a policy tool.

This article describes and evaluates the success of the March 2011 action as a representative of a post-1995 policy of "rare" coordinated interventions in troubled markets. It details the macroeconomic and financial circumstances that prompted the G-7 authorities to act and describes the immediate effect on the exchange rate level and volatility. The article also compares this action with the two most recent U.S. foreign exchange interventions in June 1998 and September 2000 and puts these three episodes in the context of G-7 historical intervention experience.

Why study intervention? Exchange rate policy is important because foreign exchange markets are large and interconnected with stock and bond markets. Disorder, or lack of two-sided liquidity in foreign exchange markets, can spill over to other asset markets.[3] Big swings in exchange rates can affect the balance sheets of banks and other financial firms. In addition, intervention per se can potentially offer important lessons for how asset markets function. To which aspects of intervention do exchange markets react? Through what channels does intervention work?

Furthermore, because exchange rates are important prices for international trade in goods and services, swings in such rates can affect real activity and international inflation rates. For example, an excessive rise in the value of the yen could impair Japanese tradable goods industries. And in the long run, excessive volatility could discourage international trade.

The next section defines foreign exchange intervention and summarizes some research on relevant issues. This is followed by a brief modern history of intervention and then a section describing the March 2011 intervention to restrain the yen. The circumstances and results of this intervention are compared with the two other most recent U.S. interventions in September 2000 and June 1998.

FOREIGN EXCHANGE INTERVENTION

This section defines intervention, explains how and why it is conducted and why it might influence exchange rates, and briefly discusses research on its effectiveness.

What Is Foreign Exchange Intervention?

Foreign exchange intervention is the practice by monetary authorities or finance ministries of buying and selling foreign currency to influence exchange rates. In the United States, for example, the U.S. Treasury and the Federal Reserve generally collaborate on foreign exchange intervention decisions, and the Federal Reserve Bank of New York usually conducts such operations on behalf of both.[4]

Central bank purchases or sales of a foreign currency change the domestic monetary base.[5] Without additional market transactions, such actions would change interest rates, exchange rates, and ultimately prices; it would simply be ordinary monetary policy conducted in the foreign exchange market instead of domestic money markets. Developed countries typically "sterilize" their foreign exchange interventions, however, which means that the central bank reverses the effects of the foreign exchange transactions on the monetary base. For example, if the New York Fed—following the instructions of the Treasury and the Federal Open Market Committee (FOMC)—sold $500 million worth of yen (purchased dollars), the U.S. monetary base would decrease by $500 million in the absence of steril-

[3] "Two-sided liquidity" means there are substantial numbers of both active buyers and sellers in a market.

[4] The May 16, 1989, FOMC meeting transcript quotes Chairman Alan Greenspan as stating, "The Treasury has the legal lead on these [intervention] decisions. We discuss it with them but the ultimate decisions are theirs" (Federal Open Market Committee, 1989a, p. 7).

[5] The monetary base is the domestic currency in circulation plus reserves of depository institutions held at the Federal Reserve Banks. Equivalently, it can be defined as domestic credit plus foreign exchange reserves.

ization. To prevent changes in domestic interest rates and prices, the New York Fed would sterilize the intervention by buying $500 million worth of U.S. government securities, which would increase commercial bank deposits with the Federal Reserve, thereby replacing the liquidity previously lost by the sale. To prevent yen-denominated short-term interest rates from falling, the Bank of Japan would need to conduct similar open market sales of yen-denominated securities to absorb the new liquidity and completely sterilize the original transaction.[6] Almost all central banks customarily target short-term interest rates, which makes such sterilization automatic. The final net effect of such a sterilized intervention would be to decrease the supply of dollar-denominated securities relative to yen-denominated securities on the market.

Why Intervene in Foreign Exchange Markets?

The "Foreign Currency Directive" of the Federal Reserve System directs intervention to "counter disorderly market conditions," in cooperation with foreign central banks, consistent with the International Monetary Fund (IMF) Articles of Agreement; Article IV, Section 1 forbids attempts to remedy balance of payments problems by manipulating exchange rates.[7] The IMF does not precisely define "disorderly market conditions"; the concept is open to interpretation. In practice, researchers such as Edison (1993) and Almekinders and Eijffinger (1996), as well as official pronouncements, support the idea that "countering disorderly market conditions" means employing intervention to resist rapid exchange rate changes that seem contrary to perceived fundamentals. This practice is often called "leaning against the wind" or described as reducing volatility in the market.

Exchange rates are important prices that influence the time path of inflation and output. An exchange rate that is significantly away from "fundamental values" can destabilize capital and trade flows that affect inflation and output. Therefore, intervention against a recent exchange rate trend is much more likely if (i) policymakers believe that an exchange rate is "misaligned"—that is, away from its fundamental value—and (ii) the recent trend is even further away from the perceived fundamental value of the exchange rate.

The idea that exchange rates can be misaligned is controversial. Fama's (1970) efficient market hypothesis suggests that asset prices always reflect fundamentals to the point where the potential excess returns do not exceed the transactions costs of acting (trading) on that information (Jensen, 1978). But it has proven very difficult to consistently link exchange rates to fundamentals in the short run. Researchers have put forward a variety of reasons—entirely consistent with rationality—to explain the persistent deviation of exchange rates from fundamentals: risk aversion, principal-agent problems, and learning and information problems. Shleifer and Vishny (1997) explore how traders are constrained by risk and principal-agent problems. Lewis (1989) and Klein and Lewis (1993) explore how learning can affect exchange rates. Whether markets are efficient or not, many policymakers believe that exchange rates can become misaligned from their fundamental values. Therefore, discussions of intervention often include the idea of "misalignment" from fundamentals or long-run equilibrium.

Although the policymaker's view of the fundamental value of the exchange is not specified and is not obvious, one can crudely calculate a long-run tendency with reference to purchasing power parity (PPP), which holds that exchange rates reflect relative price levels in the long run. The PPP-based estimate of the long-run tendency is the predicted exchange rate from a regression

6 In an environment in which short-term interest rates are essentially zero, changes in the monetary base do not have the usual effects on interest rates and prices. Short-term interest rates in both Japan and the United States were close to the zero lower bound at the time of the March intervention. Nevertheless, sterilization would be required if the central banks had an implicit target for bank reserves.

7 The "Foreign Currency Directive" is published annually in the minutes of the first FOMC meeting of the year. For an example, see www.federalreserve.gov/monetarypolicy/fomcminutes20110126.htm. The IMF's 1981 annual report, for example, states that "A member should intervene in the exchange market if necessary to counter disorderly conditions, which may be characterized inter alia by disruptive short-term movements in the exchange value of its currency" (International Monetary Fund, 1981).

of the log of that variable on a constant, a quadratic time trend, and relative log price indexes.[8]

Panel A of Figure 1 shows the JPY/USD exchange rate and its estimated long-run trend. Panel B shows the time series of U.S. interventions in the JPY/USD exchange rate; USD purchases (sales) are positive (negative). Panel C displays a scatterplot of the interventions versus the deviations of the exchange rate from its trend. The negative relation in panel C clearly shows that U.S. authorities tend to buy dollars when the price of dollars in terms of JPY is below its long-run tendency and sell dollars when the JPY price of dollars is above its long-run tendency. Contingency analysis confirms that most official U.S. intervention in the deutsche mark (DEM) or euro (EUR) and JPY markets is consistent with pushing the exchange rate toward long-run equilibrium. For example, about 61 percent of U.S. interventions are in either the top-left or lower-right quadrant in panel C, indicating a USD sale when the dollar is strong or a USD purchase when the dollar is weak, respectively. Panel C displays the two most recent U.S. interventions in the JPY market—on June 17, 1998, and March 18, 2011—as solid blue and black markers, respectively. These interventions are the two largest JPY interventions, but they are consistent with the tendency of U.S. authorities to buy (sell) dollars when the dollar is weak (strong) relative to its long-run trend.[9]

How Might Foreign Exchange Intervention Work?

Because sterilized intervention affects neither prices nor interest rates, it does not influence the exchange rate directly through these usual mechanisms. But official intervention might affect the foreign exchange market indirectly through the portfolio balance channel, the signaling channel, and/or the coordination channel.

The portfolio balance theory recognizes that sterilized intervention changes the relative supplies of bonds denominated in different currencies. If bonds from different countries are imperfect substitutes (as seems likely) and investors have only a limited appetite for the bonds from a particular country at a given rate of return, then the relative rate of return on, say, Japanese versus U.S. bonds, must depend on the relative quantities of those types of bonds. In the example in which the Fed purchases dollars/sells yen, the intervention increases the quantity of yen-denominated bonds relative to dollar-denominated bonds. Such an increase in the relative quantity of yen-denominated assets means that international investors will require a higher return on these bonds. An immediate depreciation of the yen creates a higher expected return on yen-denominated assets without changing the long-run value of the JPY/USD. Researchers are skeptical of the portfolio balance channel's importance because interventions are typically much, much too small to significantly change the relative quantities of bonds.

The signaling channel suggests that official intervention communicates (signals) information about future monetary policy. The literature on intervention has not been kind to the signaling hypothesis. Lewis (1995) and Kaminsky and Lewis (1996) found that intervention generated perverse impacts on monetary policy in their sample. Fatum and Hutchison (1999) found that intervention had no impact on federal funds futures rates. Aside from the problem that intervention seemingly does not affect expected monetary policy, the signaling story seems implausible because monetary policy and exchange rate policy are often in the hands of different institutions. If the U.S. Treasury has primary responsibility for the value of the USD, for example, then how can intervention signal anything about monetary policy, which is in the hands of the FOMC?

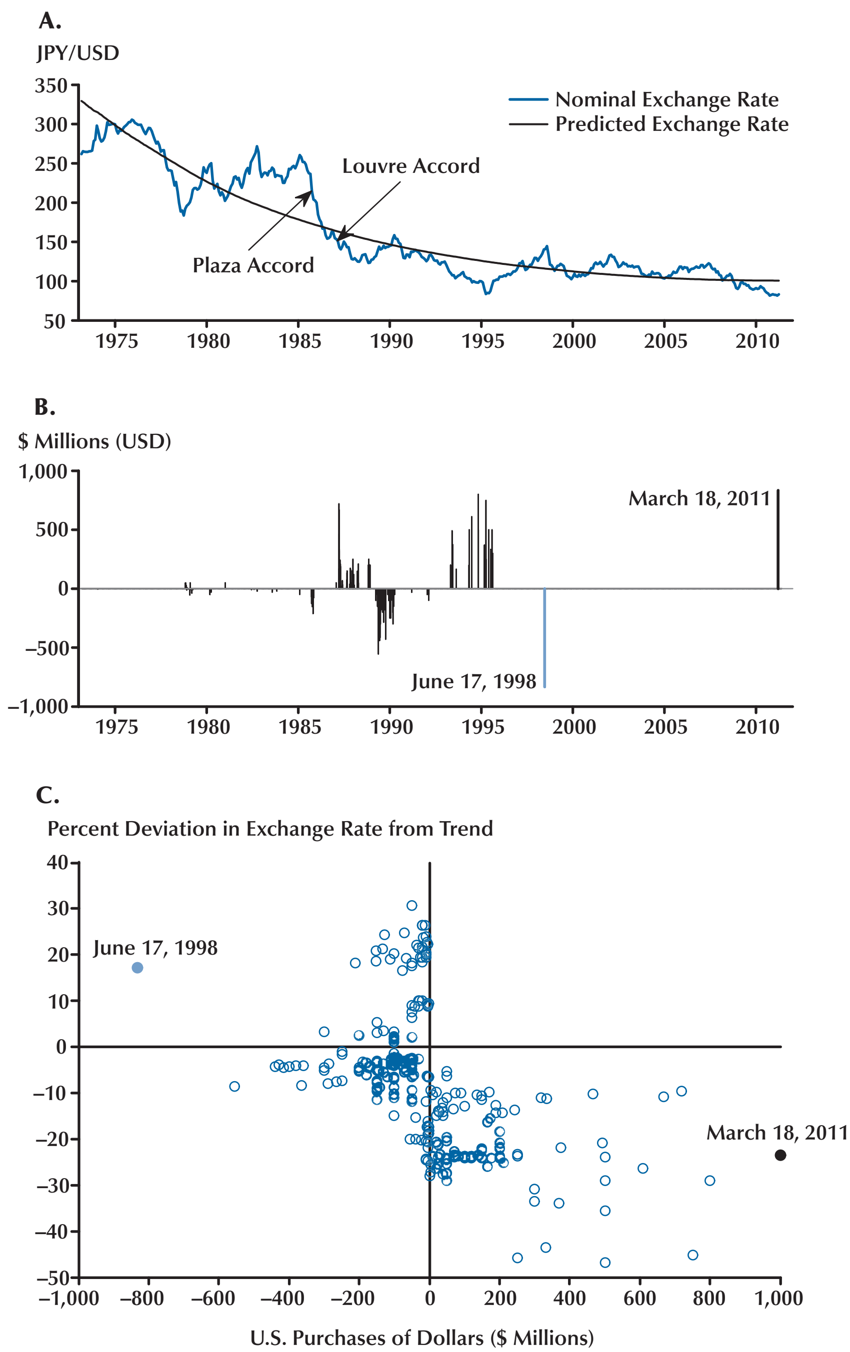
The coordination channel, which suggests that intervention might be important in coordinating the expectations of market participants, has received more attention recently. Sarno and Taylor (2001), Taylor (2005), and Reitz and Taylor (2008) have emphasized its potential importance in communicating that authorities consider that the exchange rate is deviating substantially from

[8] The quadratic time trend, which is highly statistically significant, permits a time varying Balassa-Samuelson effect, which denotes the tendency for differential productivity growth to affect real (i.e., inflation-adjusted) exchange rates.

[9] One might expect intervention sizes to grow over time with the size of financial markets.

Figure 1

The JPY/USD Exchange Rate and U.S. Intervention



NOTE: Panel A shows the JPY/USD exchange rate and a long-run equilibrium fundamental estimated from relative prices and a quadratic time trend. Panel B shows the U.S. foreign exchange intervention (millions of USD purchased) over the same sample. Panel C is a scatterplot of nonzero U.S. intervention versus the deviation from the estimated trend. The June 17, 1998, and the March 18, 2011, interventions are labeled in panels B and C.

its long-run value. During such periods, it can be extremely risky for individual investors to invest much capital in hopes of a return to long-run equilibrium. As Keynes noted, "Markets can remain irrational a lot longer than you and I can remain solvent" (see Shilling, 1993, p. 236). Nevertheless, a foreign exchange intervention can coordinate the expectations of market participants and lead investors to drive the exchange rate back toward its long-run equilibrium.

Does Intervention Work?

There is no strong consensus on the effectiveness of sterilized intervention in floating markets.[10] Sarno and Taylor (2001) report mixed evidence in their thorough survey of the literature. More recently, Bordo, Humpage, and Schwartz (2010) report on the mostly limited success of U.S. intervention efforts in the 1970s. Neely (2005) argues that the problem in finding effects of intervention is sorting out the simultaneity in the conditions under which intervention is conducted versus the effects of that intervention. Studies that consider this problem more seriously—that is, Kearns and Rigobon (2005) and Neely (2006)—find that sterilized intervention does have desired effects. Similarly, high-frequency studies, such as Fischer and Zurlinden (1999) and Payne and Vitale (2003), which tend to be less afflicted by simultaneity problems, have also found intervention to be effective. Finally, Neely (2008) surveys monetary authorities and finds that policymakers experienced in floating exchange rate markets overwhelmingly believe that intervention works in floating exchange rates.

In any case, developed countries have increasingly avoided the practice of foreign exchange intervention. Truman (2003) eloquently describes a typical view among recent policymakers:

> The evidence on the short-run effectiveness of exchange market intervention is sufficient in my view to support the judicious use of intervention by the United States as a supplementary policy instrument as long as it generally is used in a manner consistent with other economic policies; however, that same evidence falls substantially short of demonstrating that intervention is a separate policy instrument that can be used to manage exchange rates with any lasting effect.
>
> What harm is there in using an instrument that may or may not be at all effective but at least is associated about half the time with success? The harm lies in the potential for collateral damage by, for example, distracting the authorities from correcting fundamental economic policies, sending incorrect signals about those policies, or potentially moving exchange rates in directions inconsistent with those policies. These considerations suggest some of the limits on intervention as a policy tool; it may not be effective and it may not be a benign instrument (p. 248).

In other words, Truman views intervention as (perhaps) a useful tool in rare situations in which it can be used to shock or communicate with markets, but he also thinks that it cannot reliably be used to move exchange rates wherever the policymaker likes. In addition, its regular use poses the danger that policymakers might view it as a routine substitute for changes in fundamental policies.[11]

A BRIEF HISTORY OF G-7 FOREIGN EXCHANGE INTERVENTION

The coordinated intervention of March 18, 2011, is one episode in an implicit policy regime—perhaps 15 years old—that recognizes that intervention can be a useful policy tool in rare and

[10] It is important to distinguish intervention to defend fixed exchange rates, such as that conducted by the Bank of England in 1992 or the Banco de Mexico in 1994, from intervention in floating rate markets. Research is fairly clear that sterilized intervention cannot replace the use of fundamentals (i.e., monetary policy) in defending fixed exchange rates. Sterilized intervention in floating exchange rate markets is a different story, however. Unlike monetary authorities defending a fixed peg, exchange rate authorities can pick and choose the time and manner of intervention in floating rate markets.

[11] Truman was formerly the director of the Division of International Finance at the Board of Governors, so he is very familiar with the practice of intervention by the United States in the 1980s and 1990s. One might note that although the policymakers surveyed by Neely (2008) seemed to disagree with Truman's assessments of the effectiveness of intervention, they did share his concern that sterilized intervention might be used to substitute for other policies.

extreme circumstances. Prior to 1996, however, G-7 authorities intervened much more often. This section discusses the history of intervention in major currencies to explain the evolution of policy to its current state.

Pre-1973 Exchange Rate Policies

Governments have conducted policies to influence exchange rates for a very long time. Successful British, U.S., and French exchange transactions in 1927-31 to support the British pound, Austrian schilling, and German mark directly anticipated modern intervention operations (Bordo, Humpage, and Schwartz, 2007).[12] U.S. spot foreign exchange interventions to stabilize financial markets in the wake of disruptions such as the Kennedy assassination and the Cuban Missile Crisis (Bordo, Humpage, and Schwartz, 2007) presaged interventions to counter "disorderly markets" in the 1996-2011 period.

Intervention in Floating Exchange Rates (1973-1995)

After the Bretton Woods system of pegged exchange rates collapsed in March 1973, G-7 currencies floated against the dollar. Floating exchange rates tended to be more volatile than had been anticipated and neither authorities nor foreign exchange traders had much experience with them. Perhaps as a consequence, G-7 authorities tended to second-guess market movements by intervening frequently from 1973 to 1981.

Initially, the Reagan administration did not view intervention as a useful tool and intervened little from 1981 to 1985. Other countries, particularly West Germany, did intervene to attempt to stem the dollar's strong appreciation during that period. The very strong dollar and the resultant international imbalances eventually prompted the governments of France, West Germany, Japan, the United States, and the United Kingdom to sign the Plaza Accord on September 22, 1985, to cooperate to intervene to depreciate the dollar (see Figure 1). Perhaps partly as a consequence, the DEM/USD exchange rate fell 3.75 percent from September 20 to September 23 and about 37 percent between September 20, 1985, and February 2, 1987.

This large decline in the dollar's value prompted the February 22, 1987, Louvre Accord, in which Canada, France, West Germany, Japan, the United Kingdom, and the United States pledged to cooperate to stem the dollar's decline (see Figure 1). Shortly afterward, these partners intervened to purchase dollars and the dollar remained stable only for some weeks.[13]

Growing disenchantment with intervention led members of the FOMC—Governors Angell and LaWare and President Hoskins—to criticize U.S. intervention practices in 1989.[14] Kaminsky and Lewis (1996) report that, by the end of 1989, the New York Fed conducted intervention exclusively for the Treasury and no longer retained half of the intervention operations on its own books. This was very unusual as the Treasury and Federal Reserve have typically cooperated closely on such activities. It signaled, however, growing skepticism about the efficacy of intervention.

The Post-Intervention Era

By the mid-1990s, authorities of developed countries had grown skeptical about the efficacy of foreign exchange intervention operations. With only rare exceptions, the Bank of England stopped intervening after February 1993 and the Bundesbank and U.S. authorities likewise ceased the practice after 1995.[15] The Bank of Canada stopped intervening in 1998. The European

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[12] Suspicious of competitive devaluations by transactions from the British Exchange Equalization Account, the United States established the Exchange Stabilization Fund in 1934 to "stabilize" the foreign exchange value of the dollar.

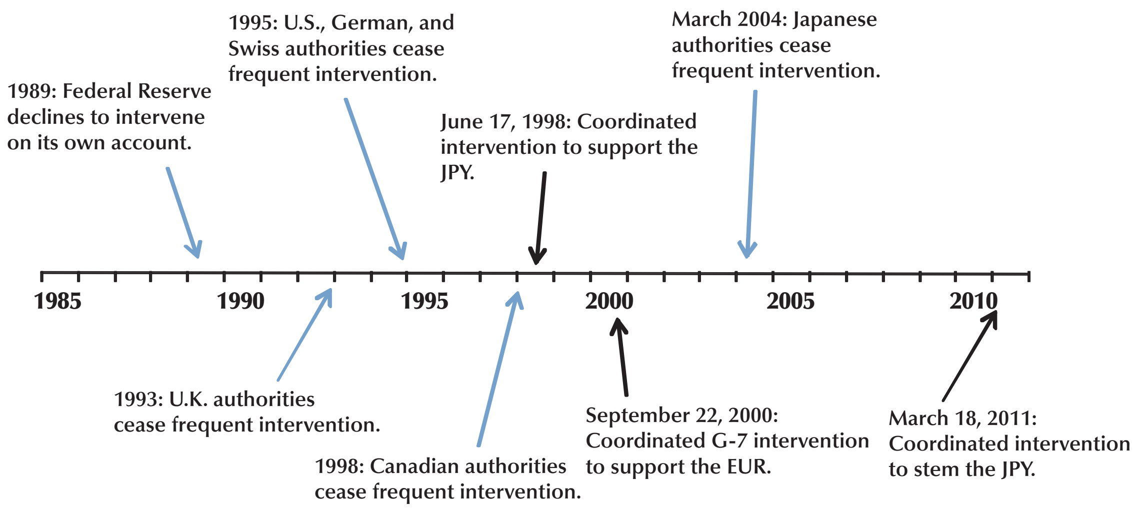
[13] Dominguez (1990) reviews U.S. intervention policies from 1985 through 1987.

[14] At the May 16, 1989, FOMC meeting, Governor Angell and President Hoskins expressed reservations about the effectiveness of intervention and suggested that the G-7 partners should change other [monetary and/or fiscal] policies. Governor LaWare dissented on a motion to authorize up to $15 billion in foreign currency balances for the Federal Reserve System (FOMC, 1989a). Criticism continued at the August 22, 1989, FOMC meeting; Governor Angell cited concerns about the consistency of a depreciating dollar with the desired goal of price stability (FOMC, 1989b).

[15] The Swiss National Bank also stopped frequent intervention in 1995. In contrast to these trends, the Reserve Bank of Australia has continued to intervene, mostly purchasing foreign currency over the past 10 years (see Reserve Bank of Australia).

Figure 2

Timeline of G-7 (and Swiss) Intervention Practices



NOTE: Timeline for G-7 (and Swiss) intervention practices after 1985.

Central Bank (ECB) has intervened only rarely since its inception in 1999.[16] These authorities have made three exceptions to this aversion to intervention: a June 1998 effort to support the JPY, a September 2000 attempt to support the EUR, and the March 2011 collaboration to restrain the JPY. Figure 2 illustrates the timeline of these actions and the three post-1995 major coordinated interventions (black arrows).

THE GREAT INTERVENTION OF 2011

The Earthquake and Its Aftermath

The economic devastation created by the March 11 earthquake was record-breaking, with hundreds of billions of dollars in property damage.[17] As might be expected in such a circumstance, Japanese markets were extremely volatile. The great uncertainty induced in Japanese financial markets increased expectations of government default and reduced equity prices. Figure 3 shows that a major Japanese equity index, the Nikkei 225, fell by about 18 percent from March 10 to March 15 and its annualized volatility skyrocketed to almost 90 percent, more than three times normal levels. Figure 4 illustrates that Japanese interest rates were already very low and did not move much.

Sound fundamental reasons existed for such financial turmoil: Figure 5 shows that the index of Japanese industrial production unexpectedly fell precipitously from February to March, from over 100 to less than 82. Simultaneously, as a result of the destruction of production facilities and the need for emergency aid, the Japanese trade balance declined from a $5 billion surplus in February to approximately zero in March and a $6 billion deficit in April (Figure 6). Reduced exports explained almost the entire decline.

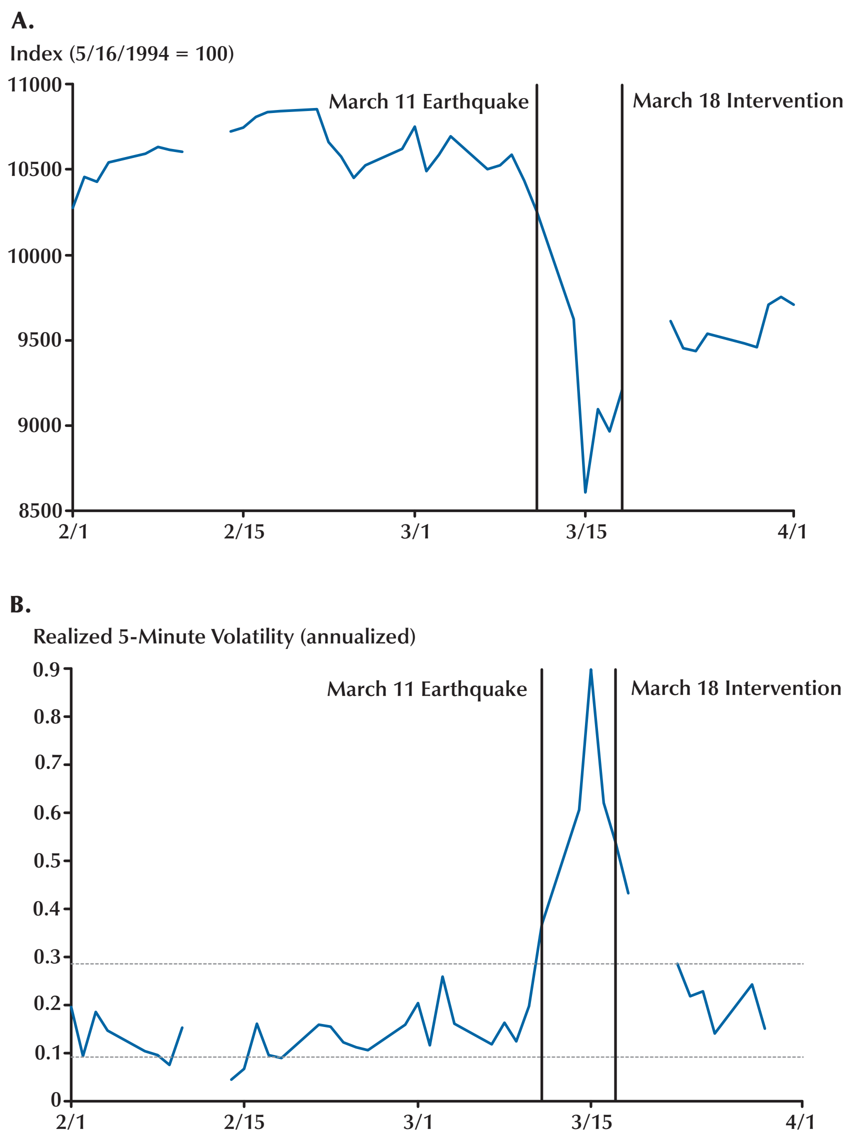
The financial press reported that two considerations dominated foreign exchange market reaction to the earthquake. First, participants expected that Japanese insurance companies

[16] Japan, traditionally very mindful of the level of the yen, continued to intervene fairly frequently until March 2004, since which it has intervened only twice.

[17] See Hosaka (2011) and National Police Agency of Japan (2011).

Figure 3

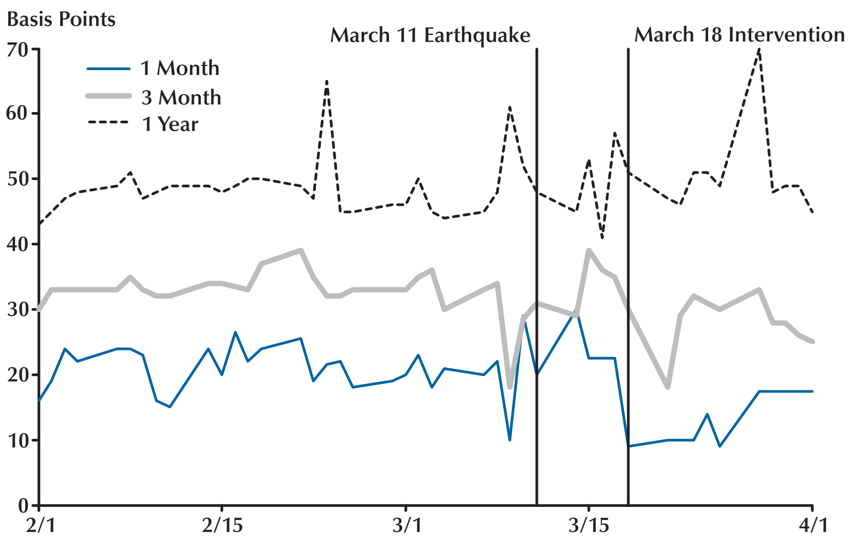
The Nikkei 225 After the Japanese Earthquake



NOTE: Panel A shows the Nikkei 225 index from February 1, 2011, through March 31, 2011; panel B shows realized volatility for the same period computed from 5-minute squared returns. The vertical lines denote the March 11 earthquake and the March 18 intervention. The horizontal lines in panel B show the 10th and 90th percentiles of volatility for the Nikkei 225 index from July 1, 2003, to March 31, 2011.

Figure 4

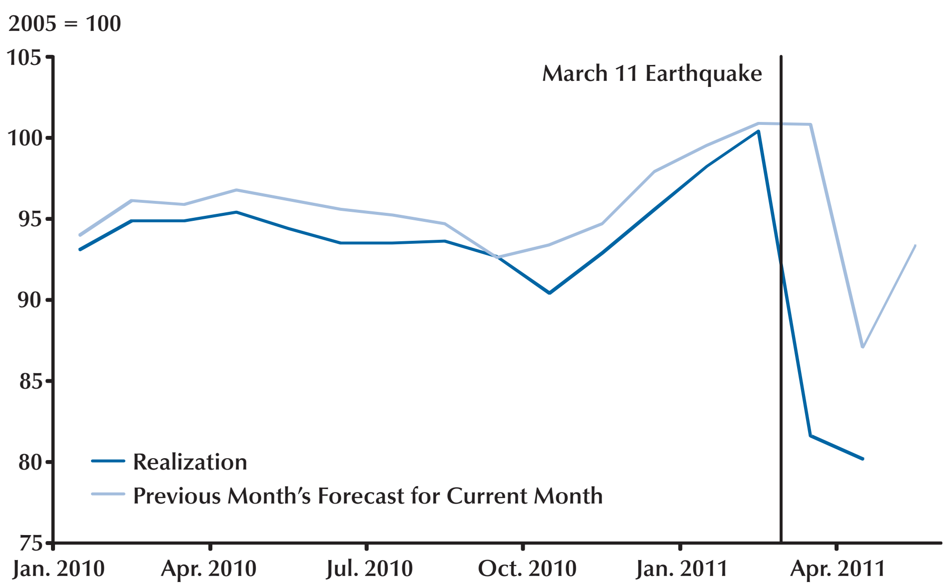
Japanese Interest Rates After the Earthquake



NOTE: The figure shows 1-month, 3-month, and 1-year Japanese interest rates from February 1, 2011, through March 31, 2011. The vertical lines denote the March 11 earthquake and the March 18 G-7 intervention.

Figure 5

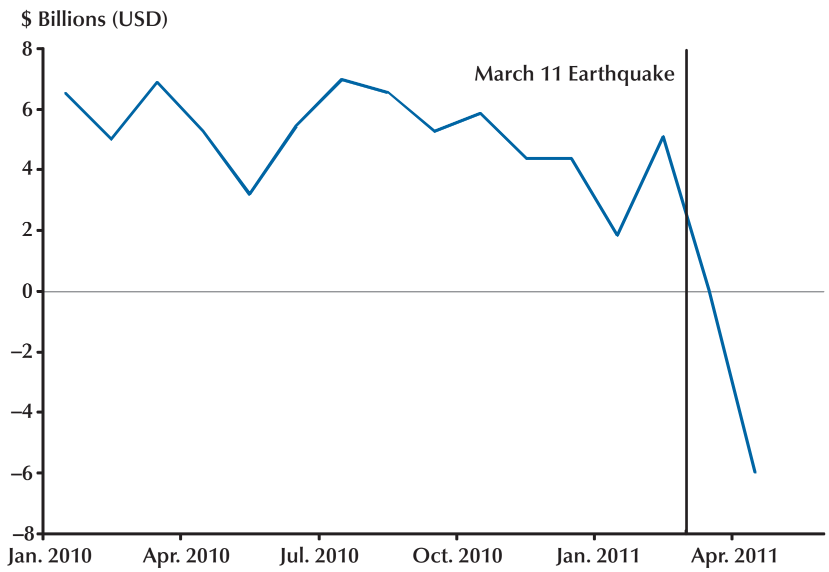
Japanese Industrial Production



NOTE: The dark blue line shows Japanese industrial production; its one-month-ahead forecast is shown by the light blue line. The vertical line denotes the break between the February and March observations.

Figure 6

Japanese Trade Balance



NOTE: The vertical line denotes the break between February and March observations.

would need to liquidate and repatriate reserves held as foreign assets. Second, Japanese investors in the carry trade—in which one borrows in low interest rate countries, such as Japan, to invest in high interest rate countries—chose to close out their Japanese borrowing in anticipation of the need for the funds. Both factors tended to strengthen the yen. The yen rapidly appreciated by 5.1 percent against the USD after the earthquake: from 82.98 JPY/USD at noon U.S. eastern time on March 10 to 78.74 JPY/USD on March 17, 2011 (Figure 7). Given an 18 percent decline in industrial production within one month and plunging exports, this rapid rise in the yen's value seriously threatened the health of Japanese tradable industries.

At the same time, JPY/USD volatility rose to unusually high levels after the earthquake. Figure 8 shows that (i) realized volatility reached an annualized level of more than 50 percent on March 16, which was three times the 90th percentile of its distribution, and (ii) 1-week implied volatility reached 22 percent on March 17.[18]

Figure 9 shows that JPY/USD trading volume grew rapidly after the March 11 quake, about tripling by March 17. G-7 officials were reportedly concerned about the effect of this turmoil on financial markets already under stress from the European debt crisis and turbulence in the Middle East. Although the earthquake clearly changed fundamentals and the appropriate path of the yen's value, Japanese officials did not think that the fundamentals justified the observed price movements.

The Intervention to Stem the Yen's Rise

Despite this evidence of "disorderly markets," financial press reports indicated that the G-7 members were unlikely to agree to intervene.[19]

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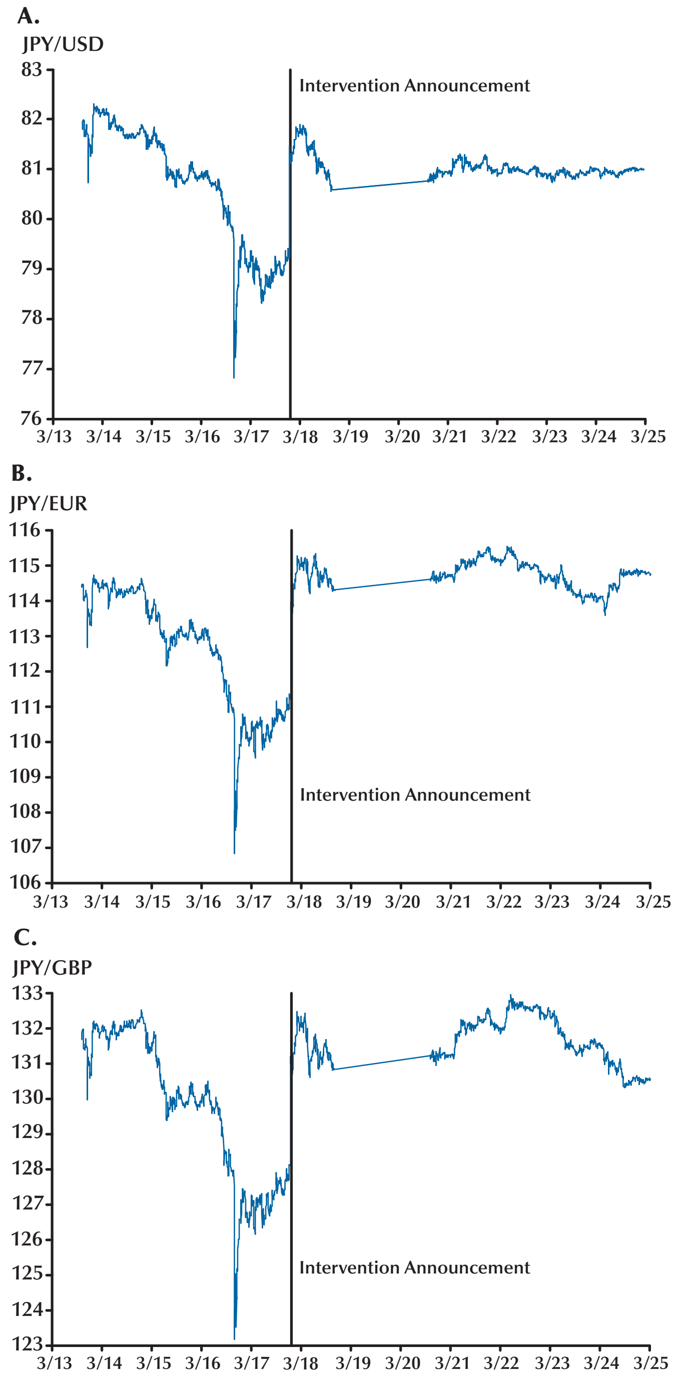
[18] Researchers have often measured volatility with either implied volatility or realized volatility. *Implied* volatility is an estimate of future volatility derived from option prices; *realized* volatility is the annualized square root of the sum of high-frequency squared returns within a day (Andersen and Bollerslev, 1998).

[19] "Yet one G7 official told Reuters that, instead, policymakers from the world's top industrialized economies are more likely to offer solidarity to Japan, the world's No. 3 economy—as opposed to agree to market intervention" (Vieira, 2011a).

Figure 7

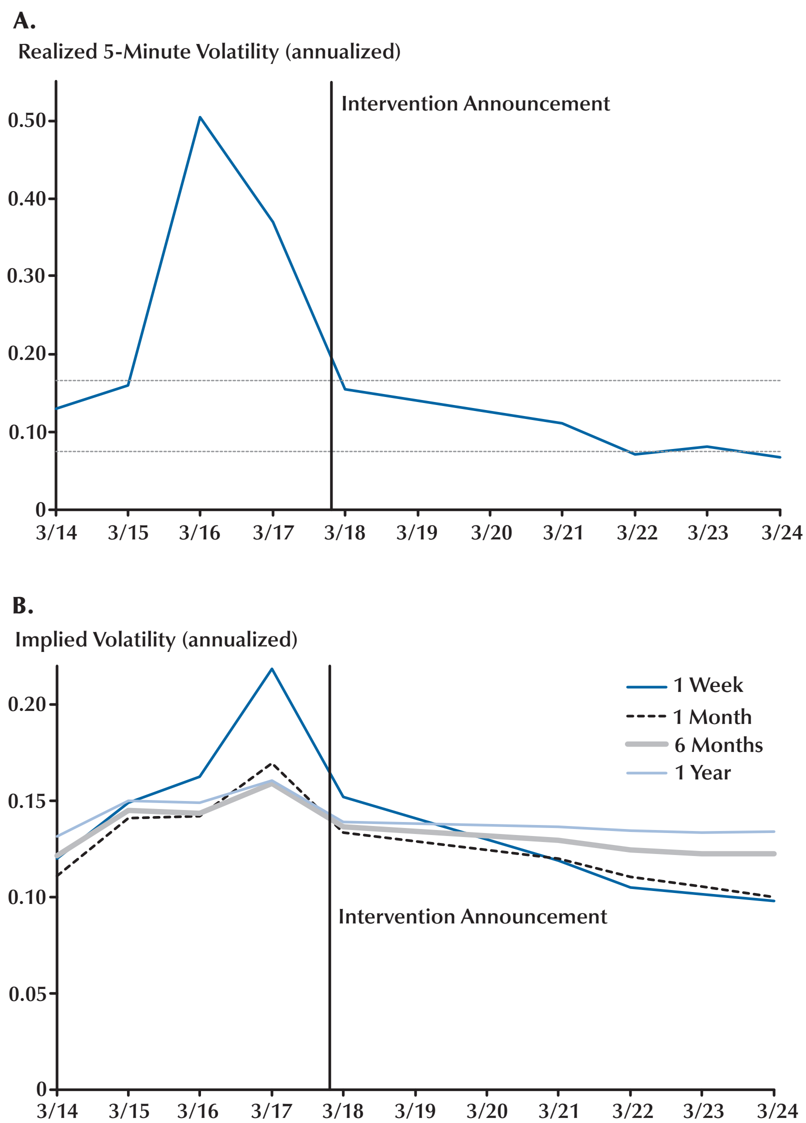
JPY Exchange Rates



NOTE: The three panels show the evolution of the JPY/USD (panel A), JPY/EUR (panel B), and JPY/GBP (panel C) exchange rates from March 13, 2011, to March 24, 2011. The vertical lines denote the date/time of the announcement of the coordinated foreign exchange intervention.

Figure 8

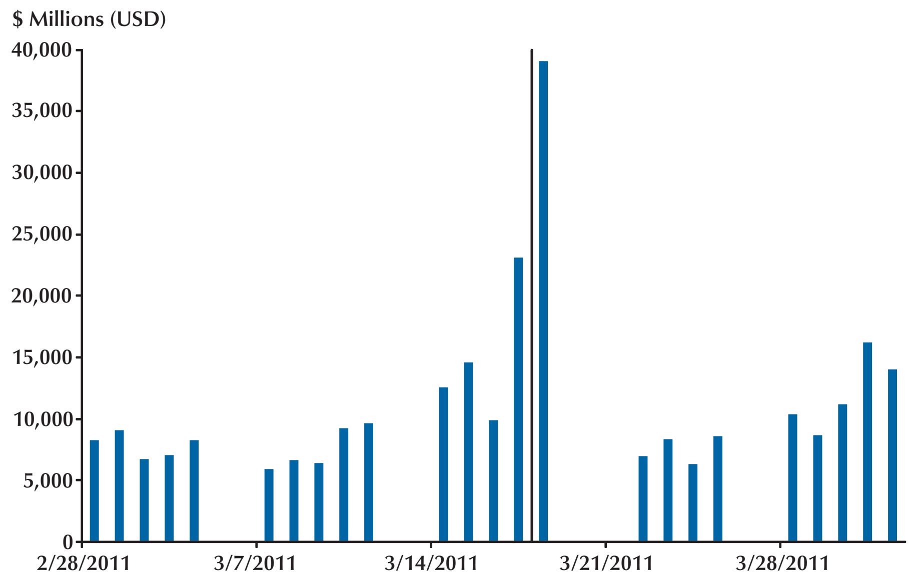
JPY/USD Realized and Implied Volatility



NOTE: Panel A shows the annualized realized volatility, computed from 5-minute squared returns, for the JPY/USD market from March 14 to March 24, 2011. The horizontal lines show the 10th and 90th percentiles of volatility for the JPY/USD over the March 26, 1998–March 31, 2011, period. Panel B shows option-implied volatility over four horizons for the same market during the same period in March 2011.

Figure 9

Daily JPY/USD Turnover in Tokyo Foreign Exchange Markets (March 2011)



NOTE: The graph illustrates the daily turnover in Tokyo foreign exchange markets during March 2011. The vertical line denotes the intervention announcement on the morning of March 18, Tokyo time.

Nevertheless, the G-7 finance ministers and central bank governors held a conference call on the evening of Thursday, March 17 (Friday morning in Tokyo) and decided to conduct a coordinated intervention to weaken the JPY. The G-7 issued a press release containing the following text:

> In response to recent movements in the exchange rate of the yen associated with the tragic events in Japan, and at the request of the Japanese authorities, the authorities of the United States, the United Kingdom, Canada, and the European Central Bank will join with Japan, on 18 March 2011, in concerted intervention in exchange markets. As we have long stated, excess volatility and disorderly movements in exchange rates have adverse implications for economic and financial stability. We will monitor exchange markets closely and will cooperate as appropriate (G-7, 2011).

Figure 7 shows that the yen reacted immediately to the intervention announcement, surging almost 4 percent within the hour against the USD, EUR, and GBP. Markets responded to the announcement of the intention to intervene rather than to the actual transactions, because efficient markets require a rapid reaction to publicly available information to preclude obvious risk-adjusted profit opportunities. One might also note in Figure 7 that the yen weakened modestly in the 18 hours preceding the announcement, perhaps in anticipation of G-7 action. Similarly, Figure 8 shows that both realized and implied volatility—both at very high levels before the intervention announcement—fell significantly the next day and even more so by Monday, March 21. Realized volatility, for example, declined below the 90th percentile of its distribution. Foreign exchange turnover declined to normal levels in the week following the intervention (see Figure 9). At the same time, Figure 3 shows that the Nikkei 225 rose about 5 percent over the weekend after the intervention and volatility declined by the end of the following week to normal levels, suggesting that financial markets viewed the intervention as favorable to Japanese corporate profits (and growth).

Are These Results Consistent with Literature Predictions?

Are the results of the March 18, 2011, intervention consistent with the best estimates in the literature on the impact of intervention? Several facts complicate such comparisons. First, the literature has typically assumed, for simplicity, that intervention has a linear impact in proportion to its size. Moreover, the fact that the clearest market reaction was to the intervention *announcement*, not to the actual intervention transactions, suggests that the intervention amount might be almost irrelevant. On the other hand, if the size of intervention is important, it might be the size relative to market turnover. Most studies of intervention have used data from the mid-1980s to the mid-1990s, but turnover in the foreign exchange market has doubled or tripled since 1995. Second, factors such as the degree of international coordination might be important but are difficult to quantify. Central bankers responding to the survey in Neely (2008), for example, tend to agree that large and coordinated interventions are more effective. It is difficult to cleanly compare the present results with those of the intervention literature because an intervention's impact might vary over time and with the nature of the exchange rate market.

Nevertheless, assuming that the intervention amount actually mattered and that markets anticipated the amounts fairly well, one can compare the observed changes in the wake of the announcement with predictions from the literature. Dominguez (2003), who studied G-3 intervention at an intraday frequency, found that a $100 million U.S. intervention in the DEM market had a maximal impact of almost 3 basis points. Using daily data, Kearns and Rigobon (2005) found that a $100 million Bank of Japan intervention had a 20-basis-point impact in the JPY/USD market. Neely (2006) found that a $100 million USD purchase caused a 5- to 6-basis-point USD appreciation in either the DEM/USD or JPY/USD markets. Extrapolating from these estimates, one might predict that a $10 billion intervention might cause a 3 to 20 percent change.

How big was the 2011 intervention? The Japanese, U.S., Canadian, and U.K. authorities announced expenditures of about $8.7 billion, $1 billion, $124 million, and $125 million, respectively, in the joint intervention. ECB authorities have not publicly released its intervention amounts, but one might interpolate from changes in the net ECB foreign currency position that it was in the neighborhood of $420 million. Therefore, the total intervention amount probably was about $10.4 billion.[20] The actual change in the exchange rate—about 4 percent—is roughly consistent with the predictions of 3.12 and 5 to 6 percent implied by Dominguez (2003) and Neely (2006), respectively, but it is considerably smaller than the 20-percentage-point estimate implied by Kearns and Rigobon (2005).

There is also a large literature that comes to mixed conclusions about the effect of intervention on volatility and higher moments (see Campa and Chang, 1998; Dominguez, 2003; Beine et al., 2007). The difficulty in separately identifying the simultaneous effects of intervention and volatility on each other might explain these mixed results. That is, intervention responds to volatility, so these variables will be positively correlated. Volatility does tend to decline in the hours and days following intervention, but it is difficult to ascertain whether the decline is the result of intervention or simply the natural tendency of very volatile markets to return to normal volatility levels over time.

Still, there is a remarkable drop in both realized (current) and implied (forward-looking) volatility associated with the March 18 intervention (see Figure 8). The fact that the short-horizon implied volatility dropped much further after the intervention suggests that the intervention did calm markets in a somewhat unexpected manner.

Comparison with June 1998

The United States has intervened in foreign exchange markets on only two other occasions since 1995: June 17, 1998, and September 22, 2000. How does the March 2011 intervention

[20] The Japanese, U.S., Canadian, U.K., and ECB intervention data are from the following sources, respectively: Ministry of Finance Japan (2011), Sack and McNeil (2011), Department of Finance Canada (2011), HM Treasury (2011), and European Central Bank (2011).

compare with those events in terms of motivation and exchange market response?

The June 1998 intervention also followed a financial crisis, the 1997 Asian exchange rate crisis in which international capital fled many developing Asian countries, such as Thailand and South Korea. In early June 1998, the main macroeconomic concern was that the yen was unusually weak and weakening further (see Figure 1), which made goods and services from other Asian countries less competitive with Japanese goods and services and harmed those countries' recoveries. Policymakers probably feared that a falling yen might cause China to devalue the renminbi (RMB), possibly sparking competitive devaluations, inflation, and instability throughout the region.

Financial markets were also in turmoil in June 1998. Panel A in Figure 10 shows that realized volatility was high in the days preceding the June 17 intervention, peaking at almost 40 percent per annum, well above the 90th percentile for its distribution. Panel B shows that the yen weakened by about 3 percent from June 11 to June 15.

The yen did strengthen modestly on June 15 and 16 in response to press reports that U.S. and Japanese officials had discussed intervention. Japanese officials had been pressing their U.S. counterparts for cooperation to raise the value of the yen. Prime Minister Hashimoto's promises of economic reform reportedly won over U.S. Treasury Secretary Rubin and President Clinton, who had been pressing for changes in Japanese policy (Dow Jones, 1998).

Rumors of strong U.S. intervention began to leak out at 8 a.m. (U.S. eastern time) on June 17 and an official statement from the U.S. Treasury confirmed the action at 8:16 a.m. (Hewett, 1998). The yen then strengthened by about 3.5 percent within the hour and realized volatility declined to normal levels by the following Monday. The intervention had its desired immediate impact.

Comparison with September 2000

The recent entry of a major new central bank to currency markets set the stage for the September 22, 2000, intervention. On January 1, 1999, the ECB began conducting a common monetary policy with a new currency, the euro, for the 11 original nations of the European Monetary Union (EMU). From its inception, the euro tended to depreciate against the dollar, falling from about 1.18 USD/EUR on the inception date to less than 0.85 USD/EUR in September 2000. Doubts about the policies of the new central bank probably contributed to this weakness. At the same time, the U.S. economy was slowing—it would officially enter a recession in March 2001—and the strong dollar/weak euro was perceived as detrimental to U.S. exporters. In addition, the Japanese feared that an overly strong yen would price Japanese exports out of the European markets (Holland, 2000). Against this backdrop, the ECB, the United States, and Japan decided to intervene to support the euro on September 22, 2000. The timing of the action was surprising; it occurred the day before a meeting of G-7 finance ministers and central bank governors.

Figure 11 shows a pattern consistent with that of the other interventions: There was a somewhat modest strengthening of the euro in the 24 hours preceding the intervention and a large (4 percent) move at the time of the intervention with perhaps a 1 percent retrenchment over the following hours. As with the other intervention episodes, volatility declined to less than the 90th percentile of its distribution within a couple of days.

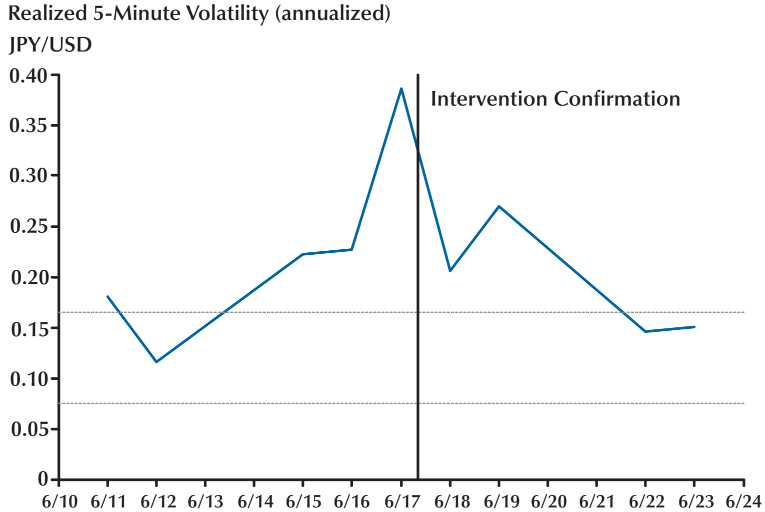
A Long-Term Effect?

A perennial question in research on foreign exchange intervention is the duration of its effects: Are the effects permanent? Unfortunately, this question cannot be answered. The nature of asset prices makes it impossible to prove that intervention has a prolonged or permanent effect, no matter what effect that intervention has. To illustrate this point, suppose that (i) the recent G-7 intervention increased the JPY/USD exchange rate by 4 percent and (ii) this effect was "permanent" in the sense that the JPY/USD rate would forever be 4 percent higher than it would have been without the intervention. Suppose, too, that JPY/USD returns have a normal distribution and an annual standard deviation (SD) of 12 percent per annum, which translates into a monthly SD of about 3.5 percent. That means that within 1 month, there

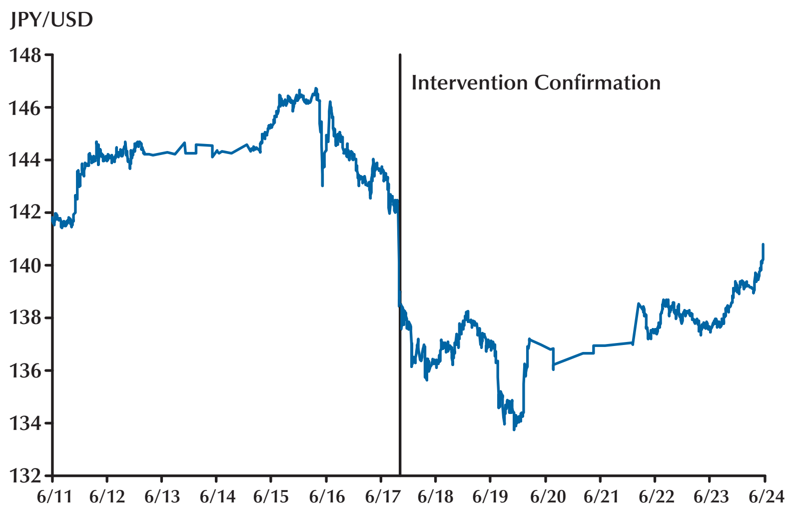
Figure 10

JPY/USD Market Behavior Near the June 17, 1998, Intervention

A.



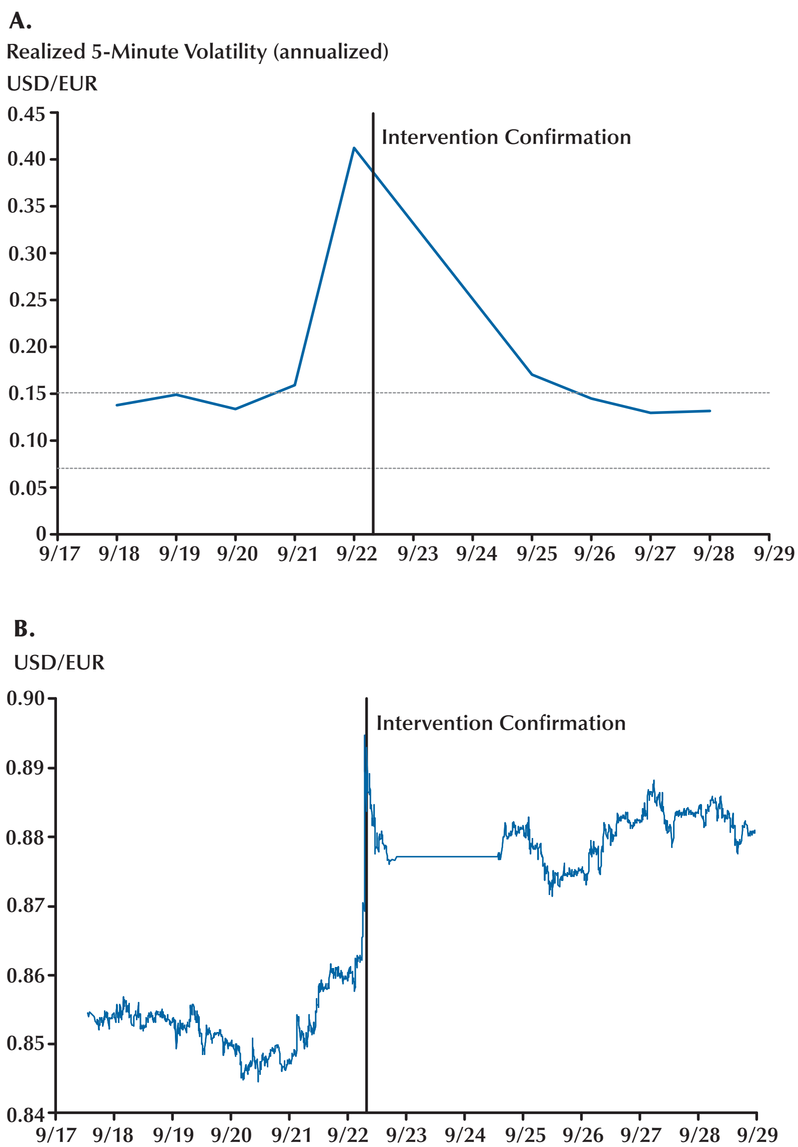
B.



NOTE: Panel A shows the daily annualized realized volatility, computed from 5-minute squared returns, for the JPY/USD market from June 11, 1998, to June 23, 1998. The horizontal lines show the 10th and 90th percentiles of volatility for the JPY/USD over the March 26, 1998–March 31, 2011 period. Panel B shows the evolution of the JPY/USD price over the same dates. In both cases, the vertical line denotes the date/time of coordinated foreign exchange intervention.

Figure 11

USD/EUR Market Behavior Near the September 22, 2000, Intervention



NOTE: Panel A shows the daily annualized realized volatility, computed from 5-minute squared returns, for the USD/EUR market from September 18, 2000, to September 28, 2000. The horizontal lines show the 10th and 90th percentiles of volatility for the USD/EUR market over the April 20, 1998–March 31, 2011 period. Panel B shows the evolution of the USD/EUR price over the same dates. In both cases, the vertical line denotes the time of coordinated foreign exchange intervention.

is a 12 percent chance that the JPY/USD rate will decline below its pre-intervention value and, within 6 months, there is a 32 percent chance that it will do so.[21] Regardless of the size of the intervention effect, the uncertainty about its effect must grow with the forecast horizon, so it can never be proved that intervention has a permanent effect.

It is also important to note that an intervention's effect need not be permanent for it to be helpful. That is, if intervention stabilizes markets by correcting misalignments or simply reintroduces a two-sided market that prevents a drastic overshoot of exchange rates, then a "permanent" intervention effect would neither be needed nor desired. Rather, the desired effect of the intervention would simply be to hasten the return to long-run equilibrium or prevent further misalignment. In either case, the long-run value of the exchange rate would be unchanged, although the intervention achieved its goal.

CONCLUSION

This article has detailed the circumstances of the March 18, 2011, intervention to weaken the Japanese yen, put current implicit intervention policy in the context of historical intervention practice, and compared the March 2011 effects with those of the two other most recent U.S. interventions: the June 1998 effort to strengthen the JPY and the September 2000 collaboration to strengthen the EUR.

After the disastrous March 11 earthquake in Japan, the yen appreciated strongly against other major currencies. The financial press cited expectations of foreign capital repatriation as the driving force behind this appreciation, which rapidly drove the yen's value above historically implied levels and increased volatility substantially, which made markets "disorderly." The G-7 finance ministers announced on the evening of March 17 that they would intervene the next day to assist Japanese authorities in stemming the yen's rise. The immediate result was a 4 percent decline in the yen's value and a large reduction in foreign exchange market volatility over the next two days. In addition, the Nikkei 225 index gained ground and its volatility returned to normal.

In several ways, the circumstances leading up to the March 18 intervention and the immediate results were similar to those of the June 1998 and September 2000 coordinated interventions. In each case, special circumstances—an earthquake, a recent financial crisis, or a new central bank and incipient recession—made exchange rate misalignments more costly than usual. Financial markets became disorderly (one-sided) as volatility increased. Press reports of intervention discussions might have caused modest exchange rate movements before the actual intervention (or announcement). Exchange rates reacted strongly and quickly to each intervention (or announcement), moving about 4 percent in the desired direction and with volatility declining substantially.

Since 1995 most advanced governments/central banks have used intervention only very sparingly as a policy tool. Examination of coordinated interventions during this period shows that intervention is not a magic wand that authorities can use to move exchange rates at will. It can be a very effective tool in certain circumstances, however, to coordinate market expectations about fundamental values of the exchange rate and calm disorderly foreign exchange markets by reintroducing two-sided risk. This article has shown that intervention remains a very effective tool, even as its use has become less common.

[21] If the annual SD of returns is 12 percent, then the annual variance of returns is $0.12^2 = 0.0144$. If returns are uncorrelated, then the 1-month and 6-month SDs of returns are

$$\sqrt{0.0144/12} \approx 0.035 \text{ and } \sqrt{0.0144/2} \approx 0.085,$$

respectively. The probabilities that the exchange rate will decline by more than 4 percent over 1 and 6 months are the values of the cumulative standard normal distribution function at –0.04/0.035 and –0.04/0.085, which are approximately 12 percent and 32 percent, respectively. In fact, the distribution of exchange rate returns has considerably fatter tails than a normal distribution, so this calculation surely understates the likelihood that the intervention will seem ineffective.

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A Comprehensive Revision of the U.S. Monetary Services (Divisia) Indexes

Richard G. Anderson and Barry E. Jones

The authors introduce a comprehensive revision of the Divisia monetary aggregates for the United States published by the Federal Reserve Bank of St. Louis, referred to as the Monetary Services Indexes (MSI). These revised MSI are available at five levels of aggregation, including a new broad level of aggregation that includes all of the assets currently reported on the Federal Reserve's H.6 statistical release. Several aspects of the new MSI differ from those previously published. One such change is that the checkable and savings deposit components of the MSI are now adjusted for the effects of retail sweep programs, beginning in 1994. Another change is that alternative MSI are provided using two alternative benchmark rates. In addition, the authors have simplified the procedure used to construct the own rate of return for small-denomination time deposits and have discontinued the previous practice of applying an implicit return to some or all demand deposits. The revised indexes begin in 1967 rather than 1960 because of data limitations.
(JEL C43, C82, E4, E50)

Federal Reserve Bank of St. Louis *Review*, September/October 2011, *93*(5), pp. 325-59.

Money is necessary to the carrying on of trade. For where money fails, men cannot buy, and trade stops.

—John Locke, *Further Considerations Concerning Raising the Value of Money* (1696, p. 319; quoted by Vickers, 1959)

Money plays a crucial role in the economy because the purchase and sale of goods and services is settled in what economists refer to as "medium of exchange." Forward-looking consumers and firms determine their desired quantities of medium of exchange at approximately the same time as they (i) form expectations of future income and expenditure and (ii) make decisions regarding desired quantities of financial and nonfinancial assets. The financial assets selected by consumers and firms may be separated into two groups. Some assets, including currency and checkable bank deposits, are innately medium of exchange—that is, usable in the purchase and sale of goods and services—while others cannot be used until converted to medium of exchange.[1] Generally, monetary assets that differ in terms of their potential usefulness as medium of exchange also differ in their own rates of return. Barnett (1980) developed the concept and theory of monetary index numbers,

[1] There are exceptions, of course. Bank checks, for example, are not accepted by all merchants. Even for currency, there are exceptions (see Twain, 1996). More seriously, currency issued by a sovereign country often is not accepted in other countries; for a discussion of monetary index numbers defined across currencies, see Barnett (2007).

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which he referred to as "Divisia monetary aggregates." Divisia aggregates measure, in a method consistent with intertemporal microeconomic theory, the aggregate flow of monetary services derived by consumers and firms from a collection of monetary assets with different characteristics and different rates of return. Underlying Divisia monetary aggregates is the concept of the user cost of a monetary asset, which is a function of the interest forgone by holding a specific asset rather than an alternative asset that does not provide any monetary services and earns a higher rate of return (referred to as the "benchmark rate"). The close connection in microeconomic theory between monetary index numbers and agents' anticipated income and expenditure suggests that monetary index numbers should be more closely related to economic activity than conventional simple sum monetary aggregates (see, for example, Hancock, 2005; Barnett and Chauvet, 2011; Barnett, forthcoming).

THE MACROECONOMICS OF MONETARY AGGREGATION

This article discusses how to construct monetary index numbers (Divisia monetary aggregates) for the United States.[2] For the most part, we do not address *when* or *why* such measurement and aggregation might be desirable, which is controversial to some extent among macroeconomists. The extant principal body of current macroeconomic analysis widely uses the concept of an aggregate measure of money and distinctly separates "money" from other assets, financial and nonfinancial.[3] Typically, macroeconomists define "money" as financial assets that either are medium of exchange or convertible to medium of exchange at de minimus cost. Demand for such assets is motivated in a macroeconomic model by either cash-in-advance or shopping-time constraints or a money-in-the-utility (or production) function specification.[4] Models differ, however, regarding whether a household or firm might replenish a depleted stock of money during the current period by selling (or using as collateral) its nonmonetary assets. If such a mechanism is permitted, the correct definition of a monetary aggregate for macroeconomic analysis depends on assumptions regarding the liquidity of those assets that are not medium of exchange.

A complementary, but alternative, line of thought argues that (i) the concept of a monetary aggregate in macroeconomics is unnecessary and misleading and (ii) models should focus on the functions of financial assets, including as a medium of exchange and an intertemporal store of value. Monetary aggregates, for example, have no role in the class of recent search-based macroeconomic models that Stephen Williamson and Randall Wright have labeled "New Monetarist economics."[5] Although the exchange of goods and services is fundamental in such models, the role of an asset as a medium of exchange is unimportant because the models (implicitly or explicitly) assume a transformation technology such that (almost) any asset can fulfill the functional role of medium of exchange—that is, all assets are liquid. For example, Williamson and Wright (2010, p. 294) write:

> Note as well that theory provides no particular rationale for adding up certain public and private liabilities (in this case currency and bank deposits), calling the sum money, and attaching some special significance to it. Indeed, there are equilibria in the model where currency and bank deposits are both used in some of the same transactions, both bear the same rate of return, and the stocks of both turn over once each period. Thus, Friedman, if he were alive, might think he had good reason to call the sum of currency and bank deposits money and proceed from there. But what the model tells us is that public and private liquidity play

[2] Throughout this analysis, the term "monetary assets" refers to those financial assets that can provide "monetary services" during the period—that is, they can serve as a medium of exchange. Some assets (currency, checkable deposits) are immediately medium of exchange. Other assets have the standby capability to act as medium of exchange if there exist markets that allow the assets to be exchanged for medium of exchange when need be, either by means of a sale or use as collateral.

[3] Walsh (2010) is a comprehensive recent textbook treatment.

[4] A classic analysis is King and Plosser (1984).

[5] Williamson and Wright (2010, 2011).

> quite different roles. In reality, many assets are used in transactions, broadly defined, including Treasury bills, mortgage-backed securities, and mutual fund shares. We see no real purpose in drawing some boundary between one set of assets and another, and calling members of one set money.

New Monetarist-style models seek to illustrate how a demand for monetary services arises as a result of optimizing behavior by households and firms. To do so, generally speaking, the models assert that a shortage of medium of exchange is costly in the sense that trades do not occur that otherwise would be Pareto welfare-improving. In such models, most financial assets are treated as near-perfect substitutes; the role of the transaction costs entailed in exchanging an asset that does not furnish medium of exchange services for one that does is secondary, such that even mortgage-backed securities furnish medium of exchange (that is, monetary) services.

In a related recent analysis that addresses neither the wisdom nor the necessity of monetary aggregation, Holmström and Tirole (2011) ask if transaction costs and "sudden stops" in financial markets explain why households and firms choose to hold larger quantities of highly liquid assets than is suggested by models with de minimus asset-market transaction costs. They note: "While some forms of equity, such as private equity, may not be readily sold at a 'fair price,' many long-term securities are traded on active organized exchanges…liquidating one's position…can be performed quickly and at low transaction costs" (p. 1). Their analysis implies that not all financial assets are perfect substitutes due to the risks that (i) market trading might suddenly halt, (ii) differential user costs can arise in the solution to the optimization problem facing households and firms, and (iii) such differential user costs reflect the differing amounts of monetary services furnished by the assets.

THE ROLE OF THE FEDERAL RESERVE BANK OF ST. LOUIS

The Federal Reserve Bank of St. Louis has published monetary index numbers (initially referred to as Divisia monetary aggregates and, later, as Monetary Services Indexes [MSI]) for two decades, beginning with Thornton and Yue (1992) and continuing with Anderson, Jones, and Nesmith (1997a,b,c) and Anderson and Buol (2005). Publication of the most recent series was suspended in March 2006 when certain necessary data became unavailable.

In this paper, we introduce a comprehensive revision of the MSI constructed at five levels of aggregation: MSI-M1, MSI-M2, MSI-M2M, MSI-MZM, and MSI-ALL. MSI-M1 and MSI-M2 are constructed, respectively, over the same components included in the Federal Reserve Board's M1 and M2 monetary aggregates. MSI-ALL is constructed over all assets currently reported on the Federal Reserve Board's H.6 statistical release (the components of M2 plus institutional money market mutual funds [MMMFs]) and is the broadest level of aggregation that currently can be constructed from available data. Finally, MSI-M2M and MSI-MZM are zero-maturity indexes (i.e., they exclude small-denomination time deposits). One change to the indexes is the adjustment of checkable and savings deposit components of the MSI for the effects of retail sweep programs, beginning in 1994.

Several changes have been made to the user costs of the components. Among these, we discontinued the previous practice of assigning an implicit return to some or all demand deposits and simplified the procedure used to construct the own rate for small-denomination time deposits. We also improved measures of savings and small time deposit rates in the Regulation Q era; as a consequence, the start date of the MSI has been changed from 1960 to 1967. Finally, the MSI are now constructed using two different benchmark rates. Our preferred benchmark rate is the maximum taken over the own rates of the components of MSI-ALL and a set of short-term money market rates (referred to in the literature as the "upper envelope") plus a small liquidity premium. The alternative benchmark rate is the larger of our preferred benchmark rate and the Baa bond yield. Previous practice had been to simply include the Baa bond yield in the upper envelope.

The remainder of the paper is organized as follows. The next section provides a brief overview of the theory behind the MSI. We then describe the MSI and their changes relative to Anderson, Jones, and Nesmith (1997c). Next, we examine the empirical properties of the MSI, emphasizing the time-series behavior of the indexes. The final section offers some conclusions.

MONETARY AGGREGATION AND INDEX NUMBER THEORY

This section briefly reviews the economic theory of monetary aggregation. Readers interested primarily in the data may skip this section without loss of continuity; readers seeking a more comprehensive survey might consult Anderson, Jones, and Nesmith (1997b).

The user cost of a monetary asset, defined as the interest income forgone by holding a specific financial asset rather than a higher-yielding asset that does not provide monetary services, plays an essential role in monetary aggregation theory. Divisia monetary aggregates are chain-weighted superlative indexes constructed over the quantities and user costs of selected sets of monetary assets. The earliest Divisia aggregates for the United States were constructed at the Federal Reserve Board through the mid-1980s by Barnett, Offenbacher, and Spindt (1981) and, later, by Farr and Johnson (1985), who introduced the descriptive label "Monetary Services Indexes."[6]

Background

Barnett (1978, 1980) developed Divisia monetary aggregates from aggregation and index number theory; see Barnett and Serletis (2000) for a comprehensive overview. The basic ideas can be illustrated with a simple money-in-the-utility function model. In each period t, a representative consumer is assumed to maximize lifetime utility:

$$\sum_{s=t}^{\infty} \beta^{s-t} u\left(\mathrm{c}_s, \mathrm{m}_s\right),$$

where c_s denotes a vector of quantities of a set of nonmonetary goods and services and m_s denotes a vector of real stocks of a set of monetary assets. The budget constraints are given by

$$\mathrm{p}_s \cdot \mathrm{c}_s = p_{s-1}^{*} b_{s-1}\left(1+R_s\right) - p_s^{*} b_s + \sum_{n=1}^{N}\left[p_{s-1}^{*} m_{n,s-1}\left(1+r_{n,s}\right) - p_s^{*} m_{n,s}\right] + Y_s$$

for all $s \geq t$, where b_s denotes the real stock of a benchmark asset that does not enter into the utility function, Y_s represents nominal income not due to asset holdings, p_s^* is a price index used to convert nominal stocks to real terms, p_s is the price vector for the nonmonetary goods and services, R_s is the nominal rate of return on the benchmark asset, and $r_{n,s}$ is the nominal own rate of return (possibly zero) for the nth monetary asset.

The user cost of each monetary asset is derived from the above maximization. Barnett (1978) derived the formula for the user cost of a monetary asset by combining individual-period budget constraints into a single lifetime budget constraint. When optimizing in period t, current-period real money balances, $m_{n,t}$, are multiplied in the lifetime budget constraint by $\pi_{n,t} = p_t^* u_{n,t}$, where

$$u_{n,t} = \frac{R_t - r_{n,t}}{1+R_t}.$$

Consequently, $\pi_{n,t}$ is the user cost for $m_{n,t}$.[7] Usually, $\pi_{n,t}$ is referred to as the "nominal user cost" and $u_{n,t}$ as the corresponding "real user cost" (Barnett, 1987, p. 118). In an alternative derivation, Donovan (1978, pp. 682-86) obtained the same expression by applying the user cost formula for a durable good to interest-bearing monetary assets.[8] Diewert (1974, p. 510) did the same for non-interest-bearing assets.

6 Divisia money measures for the United Kingdom have been maintained by the Bank of England since the early 1990s (see Fisher, Hudson, and Pradhan, 1993, and Hancock, 2005).

7 More generally, when optimizing in period t, the (discounted) user cost for $m_{n,s} (s \geq t+1)$ is given by

$$p_s^* \frac{R_s - r_{n,s}}{\left(1+R_t\right)\left(1+R_{t+1}\right)\cdots\left(1+R_s\right)}.$$

See Barnett (1978) for further discussion. Diewert (1974) provides analogous expressions for durable goods.

8 The user cost of a durable good is the difference between the purchase price of a unit of the good and the present value of the sale price one period later (adjusted for depreciation). Donovan's argument is as follows: Holding p_t^* dollars of a monetary asset in period t is equivalent to holding one real dollar of the monetary asset.

A key property in aggregation and index number theory is *weak separability*. In the present context, monetary assets are weakly separable from the other goods and services included in the utility function if

$$u(\mathrm{c},\mathrm{m}) \equiv U\left[\mathrm{c}, V(\mathrm{m})\right],$$

where U is strictly increasing in V (see Varian, 1983, p. 104). Under weak separability, utility maximization in period t implies that the vector of real money balances, m_t, chosen in that period maximizes the *sub-utility function*, $V(\mathrm{m})$, subject to the budget constraint, $\pi_t \cdot \mathrm{m} = \pi_t \cdot \mathrm{m}_t$, where π_t is a vector of nominal user costs.[9]

Chain-weighted superlative indexes constructed from data on the quantities of monetary assets and their user costs can be used to measure how $V(\mathrm{m}_t)$ evolves over time; here, we provide an overview (see the appendix for details). Specifically, the MSI are based on the superlative Törnqvist-Theil formula. The chain-weighted Törnqvist-Theil monetary quantity index is

$$V_t = V_{t-1} \prod_{n=1}^{N} \left(\frac{m_{n,t}}{m_{n,t-1}} \right)^{\frac{w_{n,t}+w_{n,t-1}}{2}},$$

where

$$w_{n,t} = \frac{\pi_{n,t} m_{n,t}}{\sum_{i=1}^{N} \pi_{i,t} m_{i,t}}$$

is the expenditure share for the nth monetary asset for period t. The index has the attractive property that its log difference is a weighted average of the log differences of its components:

$$\ln(V_t) - \ln(V_{t-1}) = \sum_{n=1}^{N} \left(\frac{w_{n,t} + w_{n,t-1}}{2} \right) \left[\ln(m_{n,t}) - \ln(m_{n,t-1}) \right].$$

Barnett (1980) interpreted the Törnqvist-Theil index as a discrete-time approximation of the continuous-time Divisia index, which is the origin of the term Divisia monetary aggregate. As he emphasized, in continuous time the Divisia index is exact for any linearly homogeneous utility function.[10]

The MSI and Their Dual User Cost Indexes

The published St. Louis MSI are constructed from nominal rather than real monetary asset quantities and, in that sense, are nominal monetary index numbers; corresponding real MSI can be obtained by dividing the nominal MSI by a price index. We also publish real user cost indexes for the various MSI that are suitable for use in empirical work as the opportunity costs of those MSI. The real user cost indexes can be multiplied by a price index to obtain corresponding nominal user cost indexes. This is analogous to the relationship between real and nominal user costs of individual monetary assets as discussed above.

Specifically, let p_t^* denote a price index, and let $M_{n,t}$ and $m_{n,t}$ denote the nominal and real quantities, respectively, of the nth monetary asset—that is, $m_{n,t} = M_{n,t}/p_t^*$. Let $u_{n,t}$ be the corresponding real user cost, which does not depend on the price index. The corresponding nominal user cost is $\pi_{n,t} = p_t^* u_{n,t}$. The published nominal MSI are constructed using nominal monetary asset quantities as follows:

$$MSI_t = MSI_{t-1} \prod_{n=1}^{N} \left(\frac{M_{n,t}}{M_{n,t-1}} \right)^{\frac{w_{n,t}+w_{n,t-1}}{2}}.$$

Thus, the purchase price of a real dollar of the monetary asset is p_t^* and the sale price of a real dollar of the asset one period later is p_{t+1}^*. If the asset earns interest, holding p_t^* dollars of the asset for one period results in $p_t^*(1 + r_{n,t})/p_{t+1}^*$ real dollars of the asset one period later. Consequently, the user cost of the monetary asset is

$$p_t^* - p_{t+1}^* \frac{p_t^*(1+r_{n,t})}{p_{t+1}^*(1+R_t)} = p_t^* \frac{R_t - r_{n,t}}{1+R_t} = \pi_{n,t}.$$

[9] Barnett (1982) emphasizes weak separability in choosing the components of a monetary aggregate. Varian (1982, 1983) derived necessary and sufficient conditions for a dataset to be consistent with utility maximization and weak separability. A number of studies have applied tests of these conditions to determine if specific groupings of monetary assets are weakly separable. For recent examples, see Jones, Dutkowsky and Elger (2005), Drake and Fleissig (2006), and Elger et al. (2008).

[10] If m_t maximizes $V(\mathrm{m})$ subject to the budget constraint $\pi_t \cdot \mathrm{m} = \pi_t \cdot \mathrm{m}_t$ for all t and $V(\mathrm{m})$ is linearly homogeneous, then in the continuous-time case

$$\frac{d\ln(V_t)}{dt} = \frac{d\ln(V(\mathrm{m}_t))}{dt} = \sum_{n=1}^{N} w_{n,t} \frac{d\ln(m_{n,t})}{dt},$$

which corresponds to the continuous-time Divisia quantity index (see Barnett, 1987, p. 141).

The expenditure shares can be computed from nominal monetary asset quantities and real user costs, since

$$w_{n,t} = \frac{\pi_{n,t} m_{n,t}}{\sum_{i=1}^{N} \pi_{i,t} m_{i,t}} = \frac{u_{n,t} m_{n,t}}{\sum_{i=1}^{N} u_{i,t} m_{i,t}} = \frac{u_{n,t} M_{n,t}}{\sum_{i=1}^{N} u_{i,t} M_{i,t}}.$$

Real MSI are obtained by dividing the nominal MSI by the price index p_t^*. This is equivalent to constructing the MSI using real quantities of individual assets when the real quantities are defined using the same price index—that is, for all n, $m_{n,t} = M_{n,t}/p_t^*$, since

$$\frac{MSI_t / p_t^*}{MSI_{t-1} / p_{t-1}^*} = \prod_{n=1}^{N} \left(\frac{M_{n,t} / p_t^*}{M_{n,t-1} / p_{t-1}^*} \right)^{\frac{w_{n,t}+w_{n,t-1}}{2}} =$$

$$\prod_{n=1}^{N} \left(\frac{m_{n,t}}{m_{n,t-1}} \right)^{\frac{w_{n,t}+w_{n,t-1}}{2}} = \frac{v_t}{v_{t-1}}.[11]$$

Let Π_t and U_t denote, respectively, the nominal and real user cost indexes for a specific MSI. The user cost index, U_t, is computed via factor reversal with its corresponding nominal MSI:

$$\frac{U_t MSI_t}{U_{t-1} MSI_{t-1}} = \frac{\sum_{n=1}^{N} u_{n,t} M_{n,t}}{\sum_{n=1}^{N} u_{n,t-1} M_{n,t-1}} =$$

$$\frac{\sum_{n=1}^{N} \pi_{n,t} M_{n,t} / p_t^*}{\sum_{n=1}^{N} \pi_{n,t-1} M_{n,t-1} / p_{t-1}^*} = \frac{\sum_{n=1}^{N} \pi_{n,t} m_{n,t}}{\sum_{n=1}^{N} \pi_{n,t-1} m_{n,t-1}}.$$

The nominal user cost index for the same MSI is $\Pi_t = p_t^* U_t$. Because nominal MSI can be converted to real MSI via division by a price index, it follows that nominal user cost indexes satisfy factor reversal with the corresponding real MSI:

$$\frac{\Pi_t V_t}{\Pi_{t-1} V_{t-1}} = \frac{\Pi_t MSI_t / p_t^*}{\Pi_{t-1} MSI_{t-1} / p_{t-1}^*} =$$

$$\frac{U_t MSI_t}{U_{t-1} MSI_{t-1}} = \frac{\sum_{n=1}^{N} \pi_{n,t} m_{n,t}}{\sum_{n=1}^{N} \pi_{n,t-1} m_{n,t-1}}.$$

CONSTRUCTING THE NEW MONETARY SERVICES INDEXES

This section describes the specification and construction of selected components of the revised St. Louis MSI. The focus is largely, but not exclusively, on aspects of the MSI that differ substantively from our earlier work (Anderson, Jones, and Nesmith, 1997c). We caution readers that this section is necessarily detail oriented, but understanding the details, though sometimes tedious, is essential if the MSI are to be used intelligently in economic research and policymaking.

Aggregation Levels, Components, and Segments

The revised St. Louis MSI introduced in this article are monthly data beginning in January 1967; when this paper was written, the most recent available data were for May 2011.[12] The indexes are published at five levels of aggregation: MSI-M1, MSI-M2, MSI-M2M, MSI-MZM, and MSI-ALL. MSI-M1 and MSI-M2 are defined over the financial assets included in the Federal Reserve Board's M1 and M2 monetary aggregates. MSI-ALL is defined over the broadest set of financial asset data currently available (to us), including the components of MSI-M2 plus institutional MMMFs. MSI-M2M and MSI-MZM are defined over zero-maturity assets—that is, financial assets immediately available for spending. More specifically, MSI-M2M is defined over the components of MSI-M2 except small-denomination time deposits, and MSI-MZM is defined over the components of MSI-M2M plus institutional MMMFs (equivalently, it includes all components of MSI-ALL except small-denomination time deposits).

Table 1 summarizes the components of the MSI. Readers should note that the number of components included in the MSI varies from month to month due to data availability. Examples of newly available data that increased the number of components include retail MMMFs (February 1973), institutional MMMFs (January 1974), other checkable deposits (OCDs) at com-

[11] This follows from the fact that the expenditure shares add up to 1.

[12] A qualification: Publication by the Board of Governors of figures regarding the deposit amounts involved in retail sweep programs lags by one month the publication of monetary data on the Board's H.6 statistical release. In constructing the MSI each month using the H.6 data, we carry forward the previous month's sweep program figures. These figures subsequently are replaced with published data as they become available.

Table 1

Components of the Monetary Services Indexes

Monetary asset	Sample period	M1	M2M	MZM	M2	ALL
Currency	Jan. 1967–present	✓	✓	✓	✓	✓
Travelers checks	Jan. 1967–present	✓	✓	✓	✓	✓
Demand deposits	Jan. 1967–present	✓	✓	✓	✓	✓
OCDs at commercial banks	Jan. 1986–present	✓	✓	✓	✓	✓
OCDs at thrift institutions	Jan. 1986–present	✓	✓	✓	✓	✓
Super NOW accounts at commercial banks	Jan. 1983–Dec. 1985	✓	✓	✓	✓	✓
Super NOW accounts at thrift institutions	Jan. 1983–Dec. 1985	✓	✓	✓	✓	✓
OCDs at commercial banks excluding Super NOW accounts	Jan. 1974–Dec. 1985	✓	✓	✓	✓	✓
OCDs at thrift institutions excluding Super NOW accounts	Jan. 1967–Dec. 1985	✓	✓	✓	✓	✓
Savings deposits at commercial banks	Sep. 1991–present		✓	✓	✓	✓
Savings deposits at thrift institutions	Sep. 1991–present		✓	✓	✓	✓
MMDAs at commercial banks	Dec. 1982–Aug. 1991		✓	✓	✓	✓
MMDAs at thrift institutions	Dec. 1982–Aug. 1991		✓	✓	✓	✓
Savings deposits at commercial banks excluding MMDAs	Jan. 1967–Aug. 1991		✓	✓	✓	✓
Savings deposits at thrift institutions excluding MMDAs	Jan. 1967–Aug. 1991		✓	✓	✓	✓
Retail money funds	Feb. 1973–present		✓	✓	✓	✓
Institutional money funds	Jan. 1974–present			✓		✓
Small-denomination time deposits at commercial banks	Jan. 1967–present				✓	✓
Small-denomination time deposits at thrift institutions	Jan. 1967–present				✓	✓

mercial banks (January 1974), money market deposit accounts (MMDAs) (December 1982), and Super negotiable order of withdrawal (Super NOW) accounts (January 1983).[13] The introduction of new assets is handled using a procedure suggested by Diewert (1980) (see Anderson, Jones, and Nesmith, 1997c, pp. 77-78, for details).

Because of data availability, each of the MSI is constructed in four segments: January 1967–December 1985, December 1985–January 1987, January 1987–August 1991, and August 1991 onward. The published MSI are created by splicing these segments at their boundaries via rescaling. Because the MSI and their segments are index numbers, the information content of the MSI is unaffected by rescaling.

The first segment and the December 1985 splice are due to the fact that we have sufficient data to treat Super NOW accounts and other OCDs as separate components from the date they enter the MSI (January 1983) through December 1985; thereafter, the MSI include the totals of Super NOW accounts and other OCDs (at both commercial banks and thrift institutions) as single components. The second segment and the January 1987 splice are due to a change in the availability of interest rate data for MMMFs as discussed below. The third segment and the August 1991 splice are due to the fact that we have sufficient data to treat MMDAs and savings deposits as separate components from the time they enter the MSI (December 1982) through August 1991; thereafter, the MSI include the total of MMDAs and savings deposits (at both commercial banks and thrift institutions) as a single component. As

[13] Throughout this article, OCDs consists of NOW and automatic transfer service accounts at depository institutions, credit union share draft accounts, and demand deposits at thrift institutions. Demand deposits are deposits at commercial banks that are legally demandable from the bank without prior notice. MMDA deposits have limited third-party transfer features as specified by the Garn-St. Germain Act of 1982 and the Dodd-Frank Act of 2010.

discussed below, the source of the data used to construct own rates for the deposit components of the MSI also differs before and after August 1991.

Retail Sweep Adjustment

Retail sweep programs at depository institutions began in January 1994. The Federal Reserve's Board of Governors described the purpose and effects of these programs in its annual report for 1999 (p. 59):

> Deposits in M1 declined further in 1998, reflecting the continued introduction of retail "sweep" programs. Growth of M1 deposits has been depressed for a number of years by these programs, which shift—or "sweep"—balances from household transactions accounts, which are subject to reserve requirements, into savings accounts, which are not. Because the funds are shifted back to transactions accounts when needed, depositors' access to their funds is not affected by these programs. However, banks benefit from the reduction in holdings of required reserves, which do not pay interest.

In the Board's H.6 statistical release, funds that have been swept from transaction deposits to savings deposits (specifically, to MMDAs) are included in published savings deposit figures rather than in transaction deposit figures. Hence, published figures for demand deposits and OCDs are too small relative to the deposit amounts that consumers and firms perceive themselves as holding at depository financial institutions, and published figures for savings deposits are too high. (Note that the published sum of transaction deposits plus savings deposits, and hence the Federal Reserve's M2 monetary aggregate, is unaffected by retail deposit sweep activity.) Precise data on the amount of deposits affected by retail deposit sweeping are not available because banks are not required to report to the Federal Reserve the amounts of deposits affected by retail sweep programs. Nevertheless, Federal Reserve Board staff estimate the amounts each month, and their estimates are available publicly on the St. Louis Fed's website.[14] Figure 1 plots demand deposits, OCDs, and their sum since 1994 both as reported on the H.6 statistical release and after adjustment for sweep effects. According to these estimates, since 2001 no less than 41 percent of total transaction deposits has been swept into savings deposits under retail sweep programs; between early 2005 and late 2008, more than 50 percent of such deposits was swept. If the estimated amounts of funds swept since 2001 were added to the Federal Reserve's M1 monetary aggregate, its level would increase by 27 percent to 36 percent. (Note: Figure 1 incorporates an estimated separation of swept amounts between demand deposits and OCDs; this allocation is based on unpublished data that are not publicly available.)

Due to insufficient data, Anderson, Jones, and Nesmith (1997c) did not consider the effects of retail sweeping on the MSI. In this revision, the MSI are constructed from data adjusted for the effects of retail sweeping (see Jones, Dutkowsky, and Elger, 2005).[15] Specifically, estimated swept amounts are added to demand deposits and OCDs and subtracted from savings deposits. The adjustment significantly affects MSI-M1 because savings deposits are not included in that index, while the effect is much smaller on the broader indexes.[16] Figure 2 plots MSI-M1 against a comparable index constructed over components not adjusted for retail sweeping; failing to adjust for the effects of retail sweeps causes significant understatement of MSI-M1.[17]

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14 See http://research.stlouisfed.org/aggreg/swdata.html. These figures are produced by the staff of the Federal Reserve Board. The precise method of estimation is not made public. For further discussion of retail sweeping, see Anderson (1995), Anderson and Rasche (2001), Dutkowsky and Cynamon (2003), Duca and VanHoose (2004), Jones, Dutkowsky, and Elger (2005), Dutkowsky, Cynamon, and Jones (2006), Elger, Jones, and Nilsson (2006), and Jones et al. (2008).

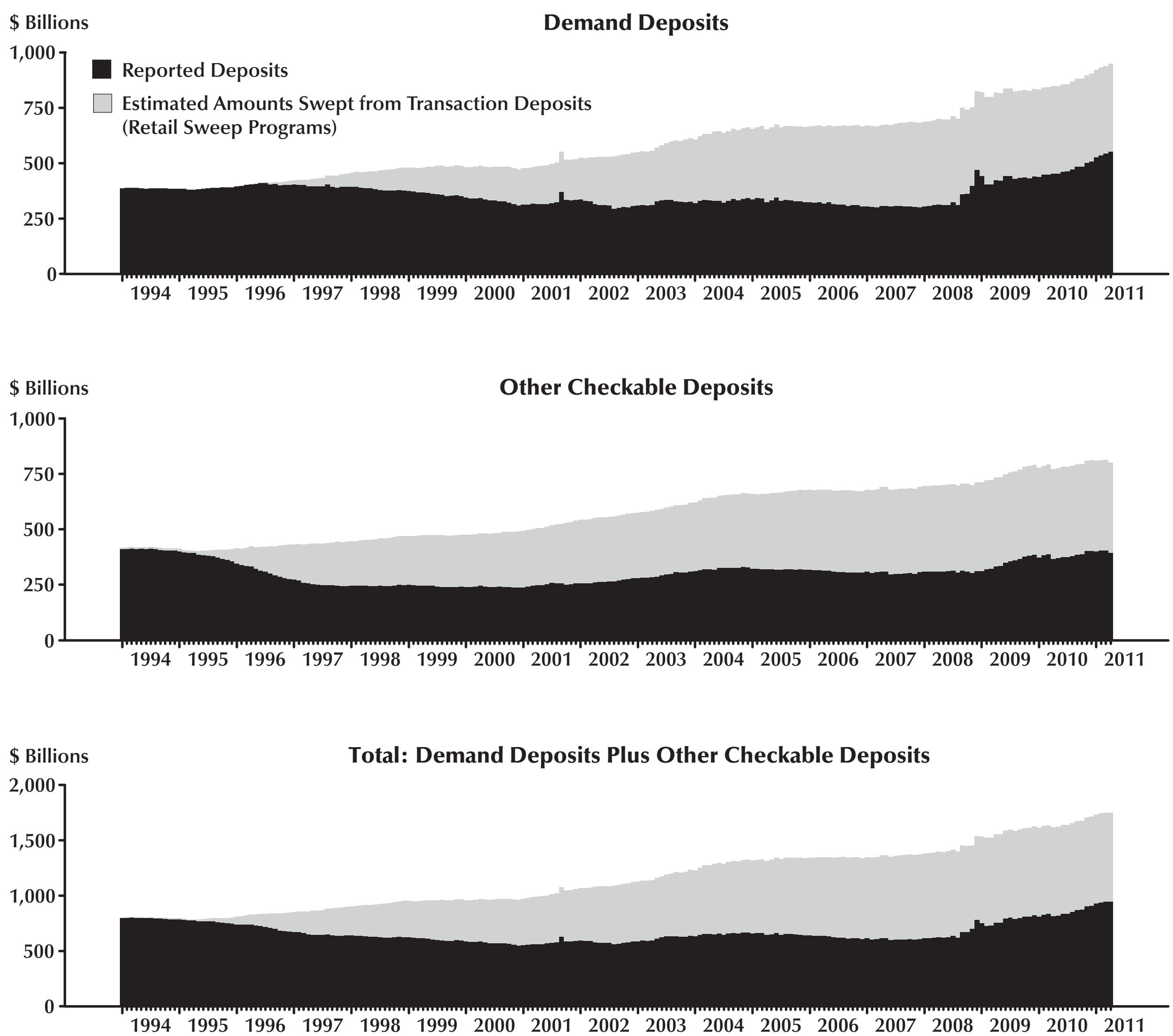
15 Jones, Dutkowsky, and Elger (2005) adjust the MSI components for the effects of both retail and commercial sweeps. We do not consider the effects of commercial sweeping in this paper. Elger, Jones, and Nilsson (2006) analyze a Divisia M1 series constructed with data adjusted for the effects of retail sweeping.

16 This, of course, also is true for summation aggregates such as those published on the Federal Reserve Board's H.6 statistical release. The Federal Reserve's M2 aggregate is unaffected by retail sweeping, while the level of the M1 aggregate is significantly reduced.

17 The indexes plotted in the figure are constructed using our preferred benchmark rate. Measurement of the benchmark rate is addressed in the next section.

Figure 1

Retail Sweeping of Transaction Deposits (1994-2011)

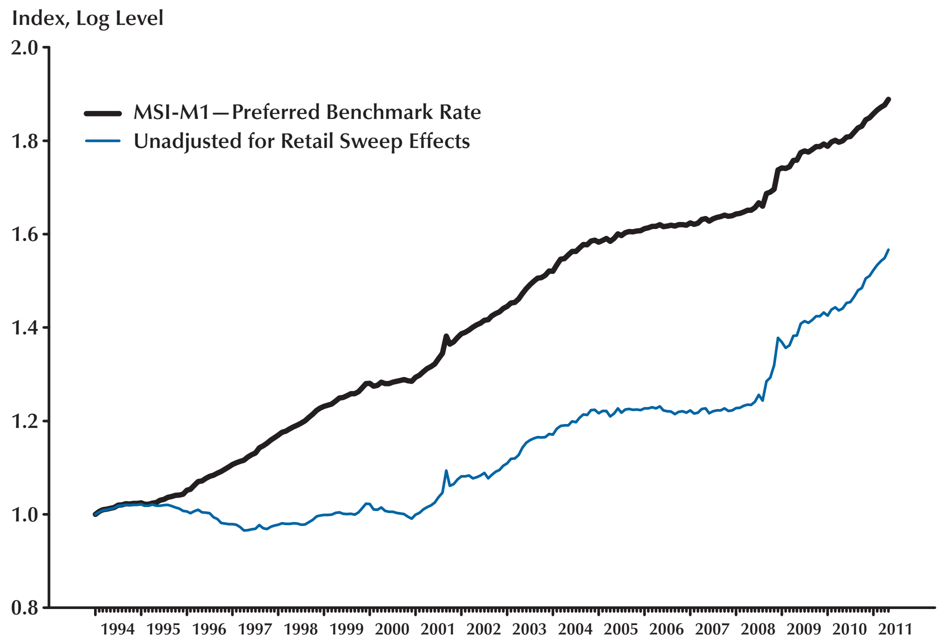


Benchmark Rates

The theory of monetary aggregation assumes that there exists a benchmark asset that furnishes no monetary services—that is, an asset that is used only to transfer wealth from period to period. It is assumed that the benchmark asset provides no "standby" or precautionary liquidity because it is not convertible into medium of exchange during the planning period at less than a confiscatory transaction cost. While the conceptual definition of the benchmark asset is straightforward, *measuring* that concept is not at all so. As noted by Fisher, Hudson, and Pradhan (1993, p. 243) and Barnett (2003, p. 50), the benchmark asset cannot be an asset that is easily traded on a secondary market. Today, however, *almost all financial assets may be used as collateral in repurchase transac-*

Figure 2

Effect of Retail Sweep Adjustments on MSI-M1 (1994-2011)



NOTE: Both indexes are scaled so that the log level is 1 in January 1994.

tions, thereby allowing conversion of the asset into medium of exchange without its sale. How, then, is the benchmark rate to be measured?

Fisher, Hudson, and Pradhan (1993) developed U.K. Divisia money measures for the Bank of England. As described by Hancock (2005, pp. 40-41), before a set of recent changes to those measures the benchmark rate was the 3-month rate on local government bills plus a premium of 200 basis points. Without the added premium, the benchmark rate would sometimes be below the own rates of some components.[18] Following Fisher, Hudson, and Pradhan (1993), Jones et al. (2008) constructed user costs for the components of MSI-M2 using the 6-month Treasury bill rate plus 200 basis points as the benchmark rate for the period from 1987 to 2004.[19]

Following Barnett and Spindt (1982), the Divisia monetary aggregates produced by the Board of Governors and subsequently by the St. Louis Fed have all taken the benchmark rate to be the maximum of the Baa corporate bond yield and the set of interest rates used to construct the component user costs. However, this approach clearly is subject to mismeasurement. In any stochastic model with forward-looking agents, the anticipated flows of income and consumption relative to expected holding-period yields (including transaction costs) determine, at least in part, agents' portfolio choices (including the quantities of monetary assets) and hence the purchased quantities of monetary services. Stracca (2004, p. 313) emphasizes the econometrician's inability to measure these expected holding period yields for long-maturity assets:

> Theoretically, the benchmark asset should be capital-certain…and at the same time provide

[18] From a measurement error perspective, if the measure of the benchmark rate generates negative user costs, then the measurement error in that measure is relatively high.

[19] Hancock (2005, p. 42) notes that a Treasury bill rate could be used as a proxy for the rate on local government bills. Bissoondeeal et al. (2010) construct a U.K. index following this suggestion.

> no liquidity services altogether. Long-term bond yields are often used as benchmark rates, but this approach is somewhat problematic…In fact, if agents have a relatively short time horizon in their portfolio allocation, what matters as an opportunity cost is not the long-term yield to maturity of the bond portfolio, but rather its expected short-term rate of return. However, this expected return cannot be observed directly, and must be proxied in some way.

Barnett (2003, p. 50) argues against including a bond yield in the calculation, preferring instead to "add to the upper envelope [over the component yield-curve-adjusted rates of return] a rate structure premium representing the premium for giving up the liquidity of the assets within the envelope."[20]

Stracca (2004) constructed a euro-zone Divisia M3 monetary aggregate along these lines. He used a short-term market interest rate to represent the own rate for the marketable securities in M3 and that rate exceeded the own rates of the other components of M3.[21] As he explains (p. 317), "We assume that the marketable instruments included in [euro] M3 provide some limited liquidity service and that they are risk free. Under these assumptions, the rate of return on a risk-free short-term financial asset providing no transaction services should be given by a short-term market interest rate plus a 'liquidity services premium.'" In practice, he set the liquidity premium equal to 60 basis points, which was the average spread between the short-term market rate and a 10-year government bond rate. He found that (i) "similar values of the premium lead to very similar patterns of the Divisia monetary aggregate" and (ii) the "annual growth rate of the Divisia index computed in this manner is very close to—indeed almost indistinguishable from—that of a Divisia index computed taking the 10-year market interest rate as the benchmark rate."

We find Stracca's (2004) reasoning compelling—namely, that the benchmark rate should exceed short-term money market interest rates. The benchmark rate also must generate positive user costs for components of MSI-ALL in order to calculate that index. As such, we begin by taking the maximum in each period over a set of rates that includes all own rates of the components of our broadest index (MSI-ALL), as well as the short-term yields on selected money market instruments, such as Treasury bills, commercial paper, eurodollar deposits, and negotiable certificates of deposit.[22] We include yields on instruments with maturities of up to six months to be consistent with our measurement of the own rate on small-denomination time deposits (discussed below). Following the literature, we refer to this maximum rate as the upper envelope. The upper envelope usually, though not always, equals one of the short-term money market yields. Following our previous discussion, we construct a benchmark rate by adding a liquidity services premium to this upper envelope. Doing so, however, requires that we determine a reasonable value for the liquidity premium. We construct MSI using a benchmark rate equal to the upper envelope plus a constant (that is, not time-varying) liquidity premium of 100 basis points, which we refer to as our "preferred" benchmark rate.

As discussed previously, Stracca (2004) selected a somewhat lower liquidity premium than we do (60 basis points). Another point of comparison is Jones et al. (2008), who defined the benchmark rate as the 6-month Treasury bill rate plus 200 basis points. Here, a benchmark rate so constructed exceeds our upper envelope throughout November 1982 to August 2008 by an average of 144 basis points, suggesting a higher liquidity premium of 144 basis points might be reasonable.[23] On the other hand, adding just 152 basis points (rather than 200 basis points) to the 6-month Treasury bill rate is sufficient to produce a benchmark rate that exceeds the upper envelope in all but two months over this period. In

[20] For another perspective on the benchmark rate, see Drake and Mills (2005, pp. 153-54).

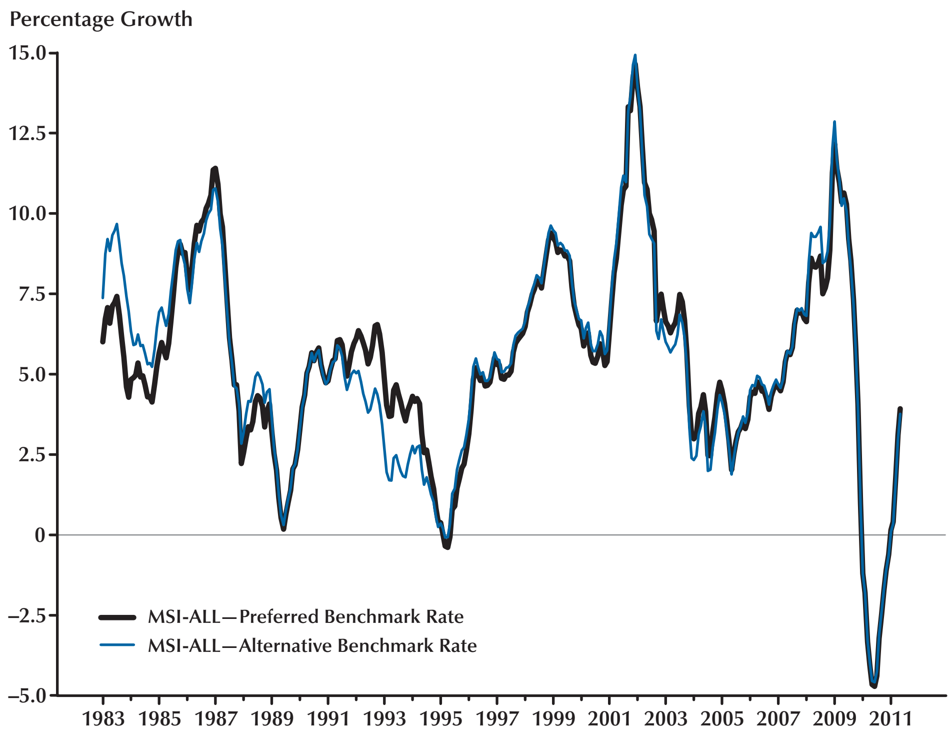
[21] See Figure 1 (p. 315) of his paper.

[22] Previously, these rates had been used to construct the MSI-L aggregate in Anderson, Jones, and Nesmith (1997c).

[23] The 6-month Treasury bill rate itself is always below the upper envelope during this period. A considerably larger premium would need to be added to the 6-month Treasury bill rate for the resulting benchmark rate to exceed the upper envelope during the height of the financial crisis. This is also the case in some earlier periods, including 1980-81.

Figure 3

Effect of Benchmark Rate on Year-over-Year Percentage Growth of MSI-ALL (1983-2011)



our judgment, adding 100 basis points to the upper envelope produces a plausible benchmark rate. Yet because this liquidity premium, essentially, is judgmental, it is important to assess the sensitivity of the resulting MSI to this choice. To measure the sensitivity of the MSI to the liquidity premium, we calculated month-to-month and year-over-year growth rates of pairs of MSI incorporating liquidity premiums of 60 and 144 basis points (rather than 100, which is used for our preferred benchmark rate) for each of the five levels of aggregation. The correlations of the growth rates for all pairs, at all five levels of aggregation, exceeded 0.98, confirming Stracca's conclusion that the behavior of monetary index numbers, including our MSI, is not highly dependent on the precise size of the liquidity premium.

We also calculate MSI using an "alternative" benchmark rate that incorporates the Baa bond yield. The alternative benchmark rate is calculated by setting the benchmark rate equal to the larger of (i) the Baa bond yield and (ii) our preferred benchmark rate, such that the alternative benchmark rate equals the Baa bond yield only in months when the bond yield exceeds the upper envelope by more than 100 basis points. Note that this calculation differs from the practice in many earlier Divisia analyses in which the Baa bond yield was included in the set of rates used to select the upper envelope and no liquidity premium was added to the upper envelope. Our rationale is that if the sum of the upper envelope plus a 100-basis-point liquidity premium (our preferred benchmark) is a sensible measure of the benchmark rate, then the Baa bond yield should be used as the benchmark rate only when it exceeds the upper envelope by more than 100 basis points. In the earlier years of our data (1967-81), our preferred benchmark exceeds the Baa bond yield approximately 58 percent of the time—that is, the preferred and alternative benchmark rates are the same more often than not. In contrast, during the later period (1982-2011), the Baa bond yield exceeds our preferred benchmark rate approximately 95 percent of the time; the exceptions are in 1989 and 2006-07.

Over the 1983-2011 period, the correlation between the growth rates of indexes calculated using the two benchmark rates (month-to-month or year-over-year) exceeds 0.93 for all five aggregation levels. Figure 3 compares year-over-year growth rates for MSI-ALL using the two benchmark rates. The comparable figure for MSI-M2 (not shown) is similar to the one for MSI-ALL. For MSI-M2M and MSI-MZM, year-over-year growth rates using the two benchmark rates differ significantly in 1983 but much less so subsequently.

Own Rates of Return

The MSI require estimates of the user costs of each component, which are derived from the spread between the benchmark rate of return and the component's own rate of return. Measurement issues with respect to the benchmark rate were discussed above. Here, we discuss issues related to own rates of return on deposits at financial institutions. With respect to measurement error, Goldfeld and Sichel (1990, p. 316) write: "A first issue concerns the own rate where there are obvious measurement difficulties created via the payment of implicit interest by the provision of services and the existence of explicit service charges. The lack of data makes it hard to evaluate the seriousness of these difficulties." Therefore, our assumptions are listed below:

Currency and travelers checks. We assume a zero own rate.

Demand deposits. We assume a zero own rate, although some demand deposits earn an implicit rate of return. As discussed by Donald Kohn (in testimony to the Committee on Banking, Housing, and Urban Affairs, U.S. Senate, June 22, 2004):

> The prohibition of interest on demand deposits distorts the pricing of transaction deposits and associated bank services. In order to compete for the liquid assets of businesses, banks have been compelled to set up complicated procedures to pay implicit interest on compensating balance accounts. Banks also spend resources—and charge fees—for sweeping the excess demand deposits of businesses into money market investments on a nightly basis.
>
> The prohibition of interest on demand deposits also distorts the pricing of other bank products. Many demand deposits are not compensating balances, and because banks cannot pay explicit interest, they often try to attract these deposits by pricing other bank services below their actual cost.
>
> Interest on demand deposits would clearly benefit small businesses, which currently earn no interest on their checking accounts. But larger firms would also benefit as direct interest payments replaced more costly sweep and compensating balance arrangements.

Barnett and Spindt (1982), Farr and Johnson (1985), and Thornton and Yue (1992) separated household and business demand deposits when constructing Divisia monetary aggregates. Household deposits were assigned an own rate of zero, but business demand deposits were assumed to earn an implicit rate of return equal to a short-term commercial paper rate net of the statutory required reserve ratio consistent with firms holding balances to compensate their banks for services used (see Mahoney, 1988, pp. 198-99 for further discussion). The data used in these studies to separate demand deposits into household and business components are not available after June 1990 (see Thornton and Yue, 1992, p. 46). Consequently, Anderson, Jones, and Nesmith (1997c) adopted an alternative procedure, assuming that the implicit rate of return on demand deposits was a fraction of the fully competitive return and applying that own rate to total demand deposits.

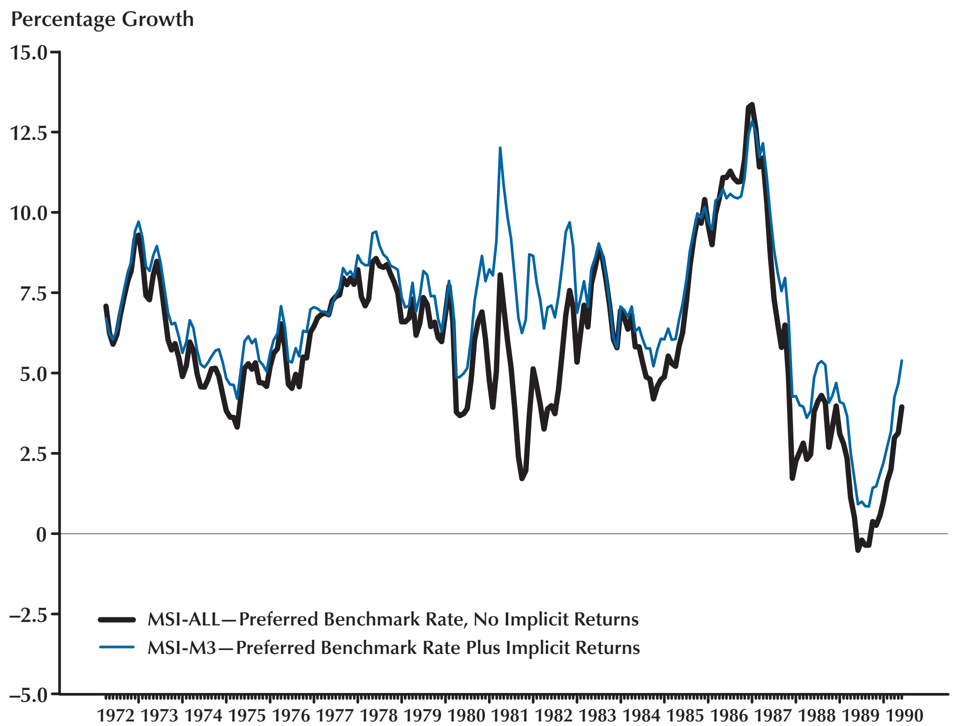
In the revised MSI reported in this article, we set the own rate of return on demand deposits to zero (rather than impute a rate of return) due to the lack of data concerning the relative ownership of demand deposits by households and businesses.[24]

[24] In addition, commercial sweeping of [business] demand deposits into Treasury bills, institutional MMMFs, eurodollar accounts, and similar liquid money market instruments increased significantly during the 1990s (see Jones, Dutkowsky, and Elger, 2005). (Recall that retail deposit sweeping consists of reclassifying checkable

Continued on next page

Figure 4

Effect of Implicit Returns on Year-over-Year Percentage Growth of MSI-M1 (April 1972–June 1990)



To assess the implications of the zero own rate, we created pairs of MSI for April 1971–June 1990 at all five levels of aggregation, in which we assigned a zero own rate to demand deposits in one MSI and, in the other assigned business demand deposits an own rate equal to the one-month commercial paper rate (adjusted for the statutory reserve requirement tax on banks) and household demand deposits an own rate of zero, following closely Farr and Johnson (1985).[25] (All pairs used our preferred benchmark rate.) The correlation of the growth rates (month-to-month and year-over-year) of the pairs exceeds 0.98 except at the M1 level of aggregation, where the correlation was 0.93 for month-to-month growth rates and 0.91 for year-over-year growth rates. Figures 4 and 5 show year-over-year growth rates for the pairs of MSI-M1 and MSI-ALL, respectively.

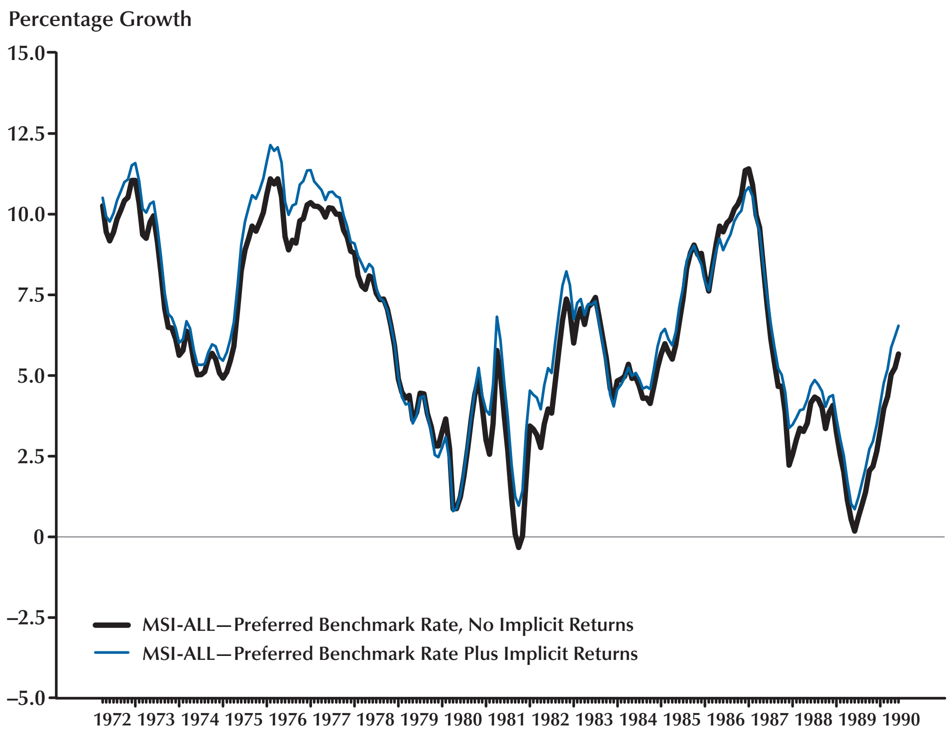
Money market mutual funds. The own rates of return for MMMFs are unpublished data obtained from the Federal Reserve Board. A single rate is available beginning in June 1974, while separate rates for retail and institutional MMMFs are available beginning in January 1987. We use the separate rates when they are available and, as mentioned above, splice the MSI in January

deposits as MMDA deposits for the purpose of calculating statutorily required reserves; these deposits do not leave the bank.) To the extent that demand deposits are swept into institutional MMMFs, they will be included in the MSI that contain those MMMFs—specifically, MSI-MZM and MSI-ALL. To the extent that the deposits are swept into money market instruments omitted from the MSI and that the owners of these swept funds continue to regard them as money, the MSI will understate the flow of monetary services received and consumed by the owners of these deposits. Of course, summation aggregates such as those reported on the Federal Reserve Board's H.6 statistical release will be understated for the same reason (see Cynamon, Dutkowsky, and Jones, 2006).

[25] The data used to construct the household and business demand deposit series are available back to June 1970 but the data for the one-month commercial paper rate are available only starting in April 1971.

Figure 5

Effect of Implicit Returns on Year-over-Year Percentage Growth of MSI-ALL (April 1972–June 1990)



1987.[26] Specifically, in the first two segments covering the periods up to January 1987, we apply the single rate to both retail and institutional MMMFs, whereas in the latter two segments covering the periods from January 1987 onward, we use separate rates.

Other bank and thrift deposits. Table 2 details the data used to measure own rates of return for the deposit components of the MSI up to August 1991. Available data include monthly figures for deposit own rates published between 1983 and 1997 by the Federal Reserve Board in a supplementary table (Monthly Survey of Selected Deposits [FR2042]) to the H.6 statistical release.[27] In addition, data beginning January 1987 have been purchased by the Federal Reserve from the Bank Rate Monitor Corporation. In our calculations, we choose to use the Federal Reserve's figures (rather than the Bank Rate Monitor data) through August 1991 because this allows us to treat MMDAs as a separate component for the longest possible time period. Although the Federal Reserve Board's figures are available through 1996, for continuity we choose to use the Bank Rate Monitor data in the final segment of the indexes (i.e., beginning in August 1991). A series of calculations confirmed that the choice has very little effect on the indexes. For small-

[26] Quantity figures for retail MMMFs begin in February 1973, so we must construct a proxy for the rate between February 1973 and May 1974. We do this via a regression against the overnight federal funds rate as in Anderson, Jones, and Nesmith (1997c, p. 66).

[27] These data are available from the Federal Reserve Board beginning in October 1983 but the published tables contain data going back to April 1983 for Super NOWs and to May 1983 for MMDAs. We used those additional observations to construct own rates after rescaling them to coincide with the data obtained from the Federal Reserve Board in October 1983. The first available observations for Super NOWs and MMDAs are applied for the first few months those assets are included in the MSI.

Table 2

OWN Rates for Deposit Components before August 1991

Monetary asset	Sample period	Own rate
Currency, travelers checks, and demand deposits	Jan. 1967–present	Zero
OCDs at commercial banks	Jan. 1986–Aug. 1991	Average rate paid, NOW accounts, commercial banks*
OCDs at thrift institutions	Jan. 1986–Aug. 1991	Average rate paid, NOW accounts, savings banks*
OCDs at commercial banks excluding Super NOW accounts	Jan. 1974–Dec. 1980	Minimum (5.0%†, average of most common rate, savings deposits‡)
	Jan. 1981–Jan. 1982	Minimum (5.25%†, average of most common rate, savings deposits‡)
	Feb. 1982–Dec. 1985	5.25%†
OCDs at thrift institutions excluding Super NOW accounts	Jan. 1967–Dec. 1973	Zero
	Jan. 1974–Dec. 1985	Same as commercial bank rate
Super NOW accounts at commercial banks	Oct. 1983–Dec. 1985	Average rate paid, Super NOW accounts, commercial banks*
Super NOW accounts at thrift institutions	Oct. 1983–Dec. 1985	Average rate paid, Super NOW accounts, savings banks*
Savings deposits at commercial banks excluding MMDAs	Jan. 1967–Jan. 1982	Average of most common rate, savings deposits‡
	Feb. 1982–Dec. 1983	5.25%†
	Jan. 1984–March 1986	5.5%†
	April 1986–Aug. 1991	Average rate paid, savings accounts, commercial banks*
Savings deposits at thrift institutions excluding MMDAs	Jan. 1967–Dec. 1969	Commercial bank rate plus 75 basis points§
	Jan. 1970–June 1973	Commercial bank rate plus 50 basis points¶
	July 1973–Jan. 1982	Commercial bank rate plus 25 basis points¶
	Feb. 1982–March 1986	5.5%†
	April 1986–Aug. 1991	Average rate paid, savings accounts, savings banks*
MMDAs at commercial banks	Oct. 1983–Aug. 1991	Average rates paid, money market deposit accounts, commercial banks*
MMDAs at thrift institutions	Oct. 1983–Aug. 1991	Average rates paid, money market deposit accounts, savings banks*

Table 2, cont'd

OWN Rates for Deposit Components before August 1991

Monetary asset	Sample period	Own rate
Small-denomination time deposits at commercial banks	Jan. 1967–March 1970	Average of most common rate, consumer-type time deposits‡
	April 1970–June 1976	Average of most common rate, time deposits in denominations of less than $100,000, maturing in less than 1 year‡
	July 1976–May 1978	Average of most common rate, time deposits in denominations of less than $100,000, other than domestic governmental units, maturing in 90 up to 180 days‡
	June 1978–Sep. 1983	Variable ceiling rate, money market time deposits, 6 months, commercial banks#
	Oct. 1983–Aug. 1991	Average rate paid, interest-bearing deposits with balances of less than $100,000 with original maturities of 92 to 182 days, commercial banks*
Small-denomination time deposits at thrift institutions	Jan. 1967–May 1978	Commercial bank rate plus 25 basis points**,††
	June 1978–Sep. 1983	Variable ceiling rate, money market time deposits, 6 months, thrifts#
	Oct. 1983–Aug. 1991	Average rate paid, interest-bearing deposits with balances of less than $100,000 with original maturities of 92 to 182 days, savings banks*

NOTE: *The data are obtained from the Federal Reserve. †See Table 8, *Annual Statistical Digest* (http://fraser.stlouisfed.org/publications/astatdig/), various issues. ‡Quarterly estimates are obtained from the Federal Reserve *Bulletin* (http://fraser.stlouisfed.org/publications/FRB/), various issues. Interpolated to obtain monthly values as described in text. §The Board of Governors established a maximum rate of 4 percent on savings deposits over this period; see Table 12.4A, *Banking and Monetary Statistics 1941-1970* (http://fraser.stlouisfed.org/publications/bms2/). Over the same period, Table S.4.12, *Federal Home Loan Bank Board Journal* (January 1974, p. 51) reports that the maximum rate payable on regular savings accounts was 4.75 percent for savings and loan associations. We therefore set the own rate on savings deposits at thrift institutions as the commercial bank rate plus 75 basis points. ¶The added basis points equal the spread between the maximum rates payable on savings deposits at savings and loan associations and mutual savings banks versus those at commercial banks from Table 8, *Annual Statistical Digest*, various issues. #Variable ceiling rates were obtained from the Federal Reserve *Bulletin*, various issues. **From January 1967 to December 1969, the Board of Governors established a maximum rate of 5 percent on single-maturity time deposits of less than $100,000, 30 days to 1 year; see Table 12.4, *Banking and Monetary Statistics 1941-1970*. Over the same period, Table S.4.12, *Federal Home Loan Bank Board Journal* (January 1974, p. 51) reports that the maximum rate payable on accounts with maturities of 6 months to 1 year was 5.25 percent for savings and loan associations. We therefore set the own rate on small-denomination time deposits at thrift institutions as the commercial bank rate plus 25 basis points. ††From January 1970 to June 1973, the added basis points equal the spread between the maximum rates payable on single-maturity time deposits of less than $100,000, 30 days to 1 year, at savings and loan associations and mutual savings banks versus those at commercial banks; from Table 8, *Annual Statistical Digest*, various issues. From July 1973 to May 1978, the added basis points equal the spread between the maximum rates payable on time accounts, 90 days to 1 year, at savings and loan associations and mutual savings banks versus those at commercial banks; from Table 8, *Annual Statistical Digest*, various issues.

denomination time deposits, the own rate is based on a 6-month maturity.

A key issue in constructing the MSI is that deposits were subject to interest rate ceilings under Regulation Q; our discussion here draws on Gilbert (1986). Generally speaking, Anderson, Jones, and Nesmith (1997c) assumed that the interest rate ceilings were binding and used them to form own rates for the components of the MSI going back to 1960. According to Gilbert (1986, p. 26), however, this was typically not the case before 1966:

> From the mid-1930s to the mid-1960s, the ceiling rates on time and savings deposits generally were above market interest rates and above the average interest rates paid on time and savings deposits by member banks. In 1957 and 1962, when market interest rates rose near or above the ceiling rates on savings deposits, these ceilings were raised…Thus, for the first 30 or so years of their existence, ceiling interest rates on time and savings deposits were above interest rates on Treasury securities in all but a few months, and average interest rates paid by member banks on all time and savings deposits were below the lowest ceiling rate in effect, the rate on savings deposits.

Moreover, as noted by Gilbert, thrift institutions were subject to interest rate ceilings only beginning in 1966 (see Ruebling, 1970).

Quarterly figures for interest rates paid by commercial banks on various types of time and savings deposits are available from January 1967 to January 1982; these figures were based on surveys by the Federal Reserve Board and the Federal Deposit Insurance Corporation (FDIC) (see Lefever, 1979). In the revised MSI presented in this article, monthly own rates for commercial bank deposits are estimated from these published quarterly figures. The start date for the MSI has been changed to January 1967 (previously, it was January 1960), reflecting the availability of the quarterly data and the fact that thrift rates were not subject to ceilings until 1966. Through January 1982, monthly own rates on savings deposits at commercial banks (excluding MMDAs) are obtained by interpolation from the published quarterly figures.[28] Due to a lack of data, the own rate on savings deposits at thrift institutions is set equal to the own rate at commercial banks plus the difference between the corresponding ceiling rates. From February 1982 until March 1986, the own rates are set at the corresponding ceilings. The Federal Reserve's monthly figures are available beginning in April 1986.

For small-denomination time deposits at commercial banks, the availability of quarterly figures changes through time. For January 1967—March 1970, we use the published rate for "consumer-type" time deposits. For April 1970–June 1976, we use the rate on deposits maturing in less than a year.[29] For July 1976–May 1978, we use the rate on deposits maturing in 90 to 180 days. Similar to savings deposits, monthly estimates of own rates on small-denomination time deposits at commercial banks are interpolated from the quarterly figures. For January 1967–May 1978, we set the own rate for small-denomination time deposits at thrift institutions equal to the own rate on deposits at commercial banks plus the difference between the corresponding interest rate ceilings.

Beginning in June 1978, commercial banks and thrift institutions were permitted to offer 6-month money market time deposits with ceiling rates tied to average auction yields on 6-month Treasury bills. Survey evidence suggests that rates paid were close to the ceilings. Consequently, we set the own rate on small-denomination time deposits at commercial banks and at thrift institutions for June 1978–September 1983 equal to the ceiling rates on 6-month money market time deposits. The Federal Reserve Board's monthly figures are available beginning in October 1983.

Due to lack of data, the own rate on OCDs at commercial banks (excluding Super NOWs) is assumed to be the smaller of the own rate on savings deposits at commercial banks and the ceiling

[28] We interpolate using cubic splines over periods when the ceiling rate is unchanged. When the ceiling rate changes, the last available estimate is used until the new ceiling is in effect. If the quarterly estimate is unchanged between two or more survey dates, we use that value for the intervening months.

[29] Note that the January 1970 figure for consumer-type time deposits reflects an increase in certain ceiling rates, but ceilings did not increase for deposits with maturities of less than one year. Thus, we actually use the October 1969 value until March 1970. The April 1976 estimate for deposits with maturities of less than one year is used for May and June 1976.

Table 3

Correlations: Growth Rates of MSI and Sum Aggregates*

	M1	M2M	M2	MZM	ALL
Correlations between MSI					
M2M	0.75				
M2	0.75	0.90			
MZM	0.69	0.95	0.89		
ALL	0.66	0.82	0.94	0.91	
M3	0.69	0.82	0.95	0.85	0.97

Correlations between sum aggregates

	M1	M2M	M2	MZM	ALL
M2M	0.53				
M2	0.49	0.68			
MZM	0.46	0.95	0.67		
ALL	0.36	0.58	0.88	0.72	
M3	0.25	0.32	0.75	0.38	0.79

Correlations between MSI and sum aggregates

	MSI					
Sums	M1	M2M	M2	MZM	ALL	M3
M1 (adjusted)	0.98	0.75	0.73	0.69	0.64	0.66
M2M	0.44	0.79	0.61	0.76	0.56	0.49
M2	0.44	0.62	0.78	0.63	0.75	0.71
MZM	0.38	0.74	0.59	0.80	0.64	0.53
ALL	0.32	0.52	0.68	0.67	0.80	0.73
M3	0.22	0.31	0.51	0.36	0.56	0.69

NOTE: Sample period for calculations excluding M3, January 1967–May 2011; for calculations including M3, January 1967–February 2006.
*Change from previous month, percent annual rate.

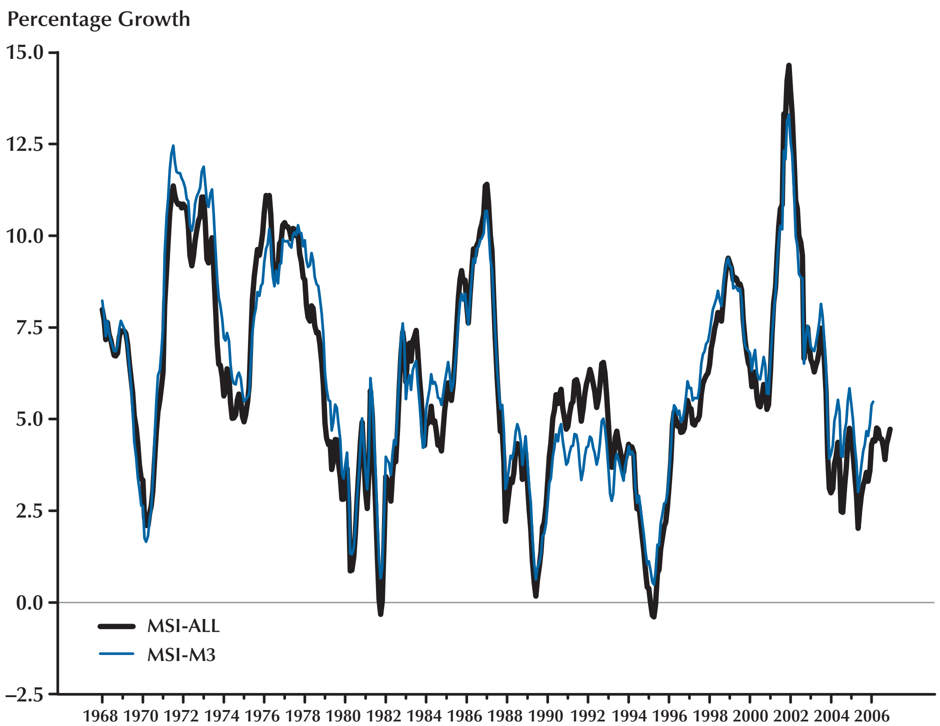
rate on NOW accounts at commercial banks from January 1974 until January 1982. For February 1982–December 1985, the ceiling rate on NOW accounts is used. The own rate for OCDs at thrift institutions is set to zero before January 1974. For January 1974–December 1985, it is equal to the own rate on OCDs at commercial banks.

Yield-Curve Adjustment of Small-Denomination Time Deposits

The stock of small-denomination time deposits includes deposits with a range of different maturities. Farr and Johnson (1985) referred to such components as "composite asset stocks." They constructed own rates for composite asset stocks by *yield-curve adjusting* the available interest rate data. The yield-curve adjustment consisted of taking an interest rate for a particular maturity and subtracting the corresponding term premium obtained from the Treasury yield curve. The stated purpose of the adjustment was as follows: "Given typical yield-curve relationships, liquidity premiums keep rates on long-maturity assets higher than those on short-maturity assets. Thus, before the rates can be compared, they must

Figure 6

Comparison of Year-over-Year Percentage Growth in MSI-ALL and MSI-M3 (1968-2006)



NOTE: Both measures are based on the authors' preferred benchmark rate.

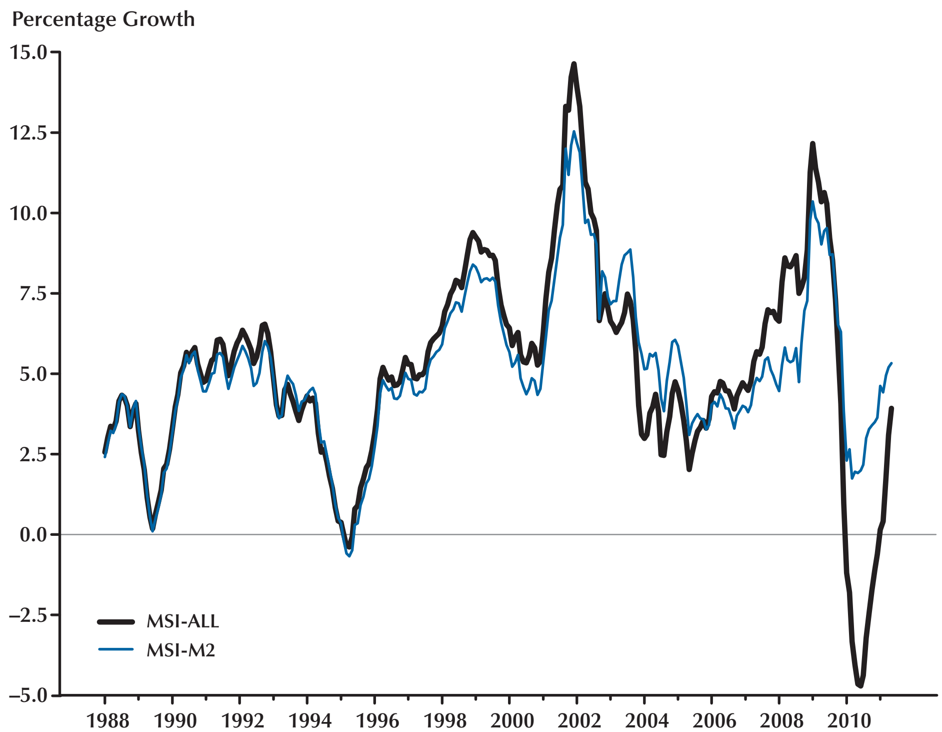
The Discontinuance of M3

The Federal Reserve Board discontinued publication of the M3 monetary aggregate in early 2006 because "M3 does not appear to convey any additional information about economic activity that is not already embodied in M2 and has not played a role in the monetary policy process for many years" (H.6 statistical release for March 23, 2006). Data on the non-M2 components of M3 were discontinued at the same time except for institutional MMMFs. Subsequently, some analysts suggested that the Board's assertion regarding the lack of additional information in M3 was in error. To assess the importance of the end of M3, we constructed an MSI-M3 through February 2006 (the most recent feasible date) and compared it with the other five MSI using our preferred benchmark rate. The top panel of Table 3 shows correlations between month-to-month growth rates of the five MSI and MSI-M3. As the table shows, MSI-M3 is most highly correlated with MSI-M2 and MSI-ALL (0.95 and 0.97, respectively). This is also the case for year-over-year growth rates (Figure 6 shows year-over-year growth rates of MSI-ALL and MSI-M3). This lends support to the view that little may have been lost, at least as far as the MSI are concerned, by the discontinuance of M3. On the other hand, it is interesting to note that year-over-year growth of MSI-M2 and MSI-ALL diverged much more than usual in 2010 due to sharp declines in institutional MMMFs (see the discussion in the main text). Figure 7 compares year-over-year growth for MSI-M2 and MSI-ALL for 1988-2011.[30] Thus, it seems reasonable to infer that an MSI-M3, if it could be constructed for the same period, would display a similar divergence, suggesting that some information value, at least during periods of financial market upheaval, was lost with the discontinuance of M3.

[30] The indexes are identical until institutional MMMFs enter in 1974.

Figure 7

Comparison of Year-over-Year Percentage Growth in MSI-M2 and MSI-ALL (1988-2011)



NOTE: Both measures are based on the authors' preferred benchmark rate.

be put on some common basis—the liquidity premiums must be extracted" (Farr and Johnson, 1985, p. 6). They then constructed user costs by using the maximum of the yield-curve-adjusted interest rates as the own rate. Anderson, Jones, and Nesmith (1997c) used a different procedure to calculate user costs of composite asset stocks but their procedure continued to involve the use of yield-curve-adjusted interest rates (see p. 76 of their paper for details).

As of this revision, we have simplified the calculation of the own rates of return on small-denomination time deposits. As discussed above, we now use the rates of return on 6-month deposits going back to the mid-1970s as the own rates. We do not yield-curve adjust any rates of return, including the short-term yields on the money market instruments included in the upper envelope. To assess the significance of this change, we constructed alternative indexes over the period from August 1991 onward that incorporated yield-curve adjustment. Specifically, we constructed alternative own rates of return on small-denomination time deposits defined as the maximum of the yield-curve-adjusted rates on the available maturities and we yield-curve adjusted all short-term yields included in the upper envelope. These indexes differed very little from our indexes without yield-curve adjustment.

EMPIRICAL ANALYSIS

This section presents an empirical analysis of our newly revised MSI. Figures 8 and 9 illustrate the short- and long-run growth of the MSI. Figures 10 and 11 show the interaction between the scope of the indexes and the aggregation method. We analyze the components of the MSI during the recent financial crisis (2007-09) in Figures 12, 13, and 14. Throughout this section, all figures are based on MSI computed with our preferred benchmark rate.

Long-Run and Short-Run Growth

Figure 8 shows both month-to-month and year-over-year MSI growth. Generally speaking, movements in the five MSI are similar. During the most recent five years, growth accelerated during 2007 and 2008 in response to Federal Reserve policies, slowed during 2009 (with negative year-over-year growth for MSI-MZM and MSI-ALL due to runoffs in institutional-type MMMFs), and strengthened during 2010 in response to expansionary Federal Reserve policies. The MSI cluster into three groups: MSI-M1, MSI-M2 and MSI-ALL, and MSI-MZM and MSI-M2M.

Figure 9 depicts growth of MSI during four selected decade-long periods. Panel A displays the late 1960s to mid-1970s, a period of increasing inflation. MSI growth slowed during 1969 as Federal Open Market Committee (FOMC) policy tightened, with decreases during 1970 in the levels of MSI-M2M and MSI-MZM. Steady growth resumed late in 1970 as the FOMC eased policy. Panel B includes the 1979-82 period of disinflation and the subsequent recovery. Growth of MSI-M1 slowed little during the period, while growth of both MSI-M2M and MSI-MZM fell rapidly beginning in mid-1978 and remained near zero until the mid-1982 easing of Fed policy. Panel C includes the 1990 recession and its subsequent "jobless recovery" plus the productivity acceleration that started in late 1992. The effects of the Federal Reserve's policy easing in 1991 and policy tightening in 1994 are apparent. Panel D includes the 2001 recession/recovery, the subsequent housing boom and financial crisis, and the Federal Reserve's credit-easing policies during 2008 and its 2009-11 quantitative easing policies. During 2009, two of the indexes, MSI-MZM and MSI-ALL, display absolute decreases due to runoffs in institutional-type MMMFs.

Method of Aggregation Versus Scope of the Aggregate

This section compares and contrasts the MSI with each other and monetary aggregates constructed by summation of the dollar amounts of the included assets; the latter are denoted as "SUM-M1" and so on. SUM-M2M, SUM-MZM, and SUM-M2 are identical to the monetary aggregates available through FRED.[31] SUM-M1 is not the same as the Federal Reserve's M1 aggregate because it is retail-sweep adjusted to be comparable with MSI-M1.[32] SUM-ALL is identical to SUM-M2 plus institutional MMMFs, which are also available through FRED.

Measurement of a monetary aggregate involves two general concepts: the *method* of aggregation (as an economic index number or via summation of the included assets) and the *scope* of the aggregate (i.e., which assets are included in the aggregate). With respect to the former, monetary aggregates produced by the Federal Reserve's Board of Governors are summation aggregates: Each is the simple unweighted sum of the dollar values of a selected set of financial assets. The MSI, in contrast, are chain-weighted superlative statistical index numbers. With respect to the scope, both the MSI and summation aggregates are (almost) nested: M1 is a proper subset of M2M and M2M is a proper subset of M2. The aggregation-level MZM differs from the level M2M by the inclusion in the former of institutional-type MMMFs. All the narrower indexes are proper subsets of the aggregate "ALL." From the standpoint of monetary aggregation/index number theory, the two issues are related since superlative index numbers should be constructed over groups of monetary assets that are weakly separable. For further discussion of this view, see Barnett (1982), Swofford and Whitney (1991), and Belongia (1996).

Figure 10 shows a scatterplot matrix of month-to-month percentage growth rates of the five MSI; correlations between the MSI are shown in the

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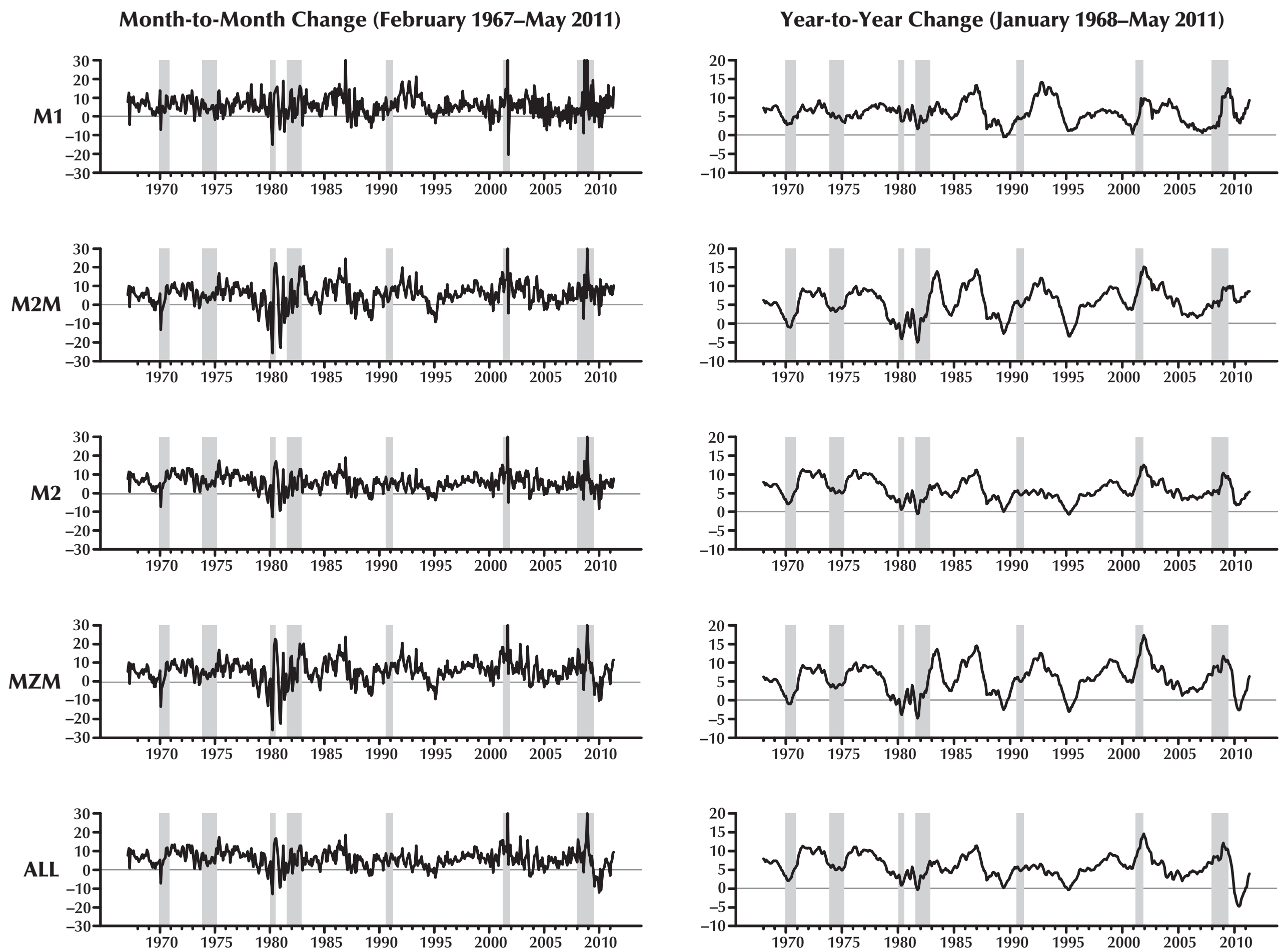
[31] FRED (Federal Reserve Economic Data); available through the St. Louis Fed's website (http://research.stlouisfed.org/fred2/).

[32] SUM-M1, as measured here, equals the Federal Reserve's M1 aggregate plus the estimated amount of retail sweeps; Dutkowsky, Cynamon, and Jones (2006) and Cynamon, Dutkowsky, and Jones (2006) refer to this as M1RS (M1 plus retail sweeps).

Figure 8

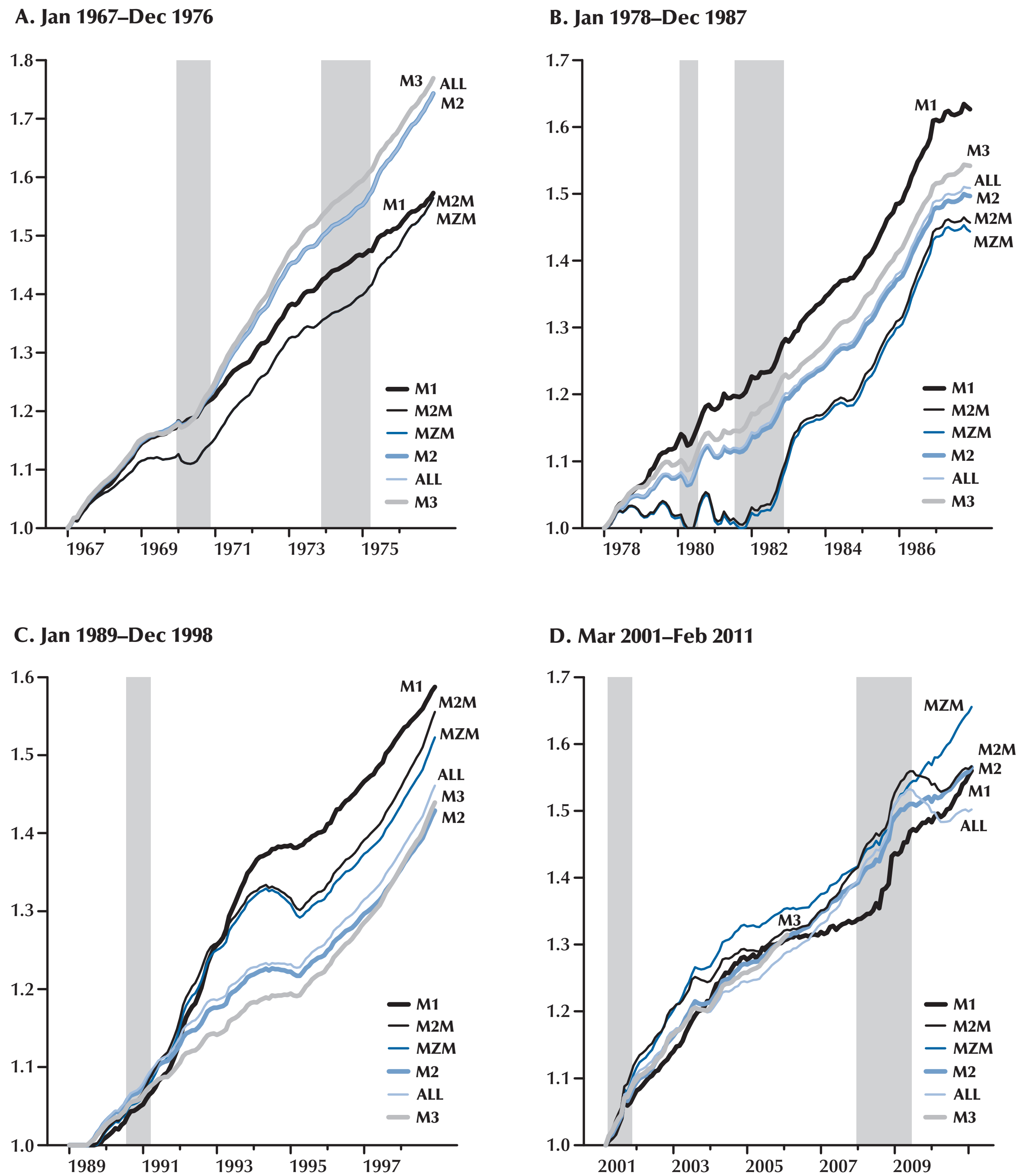
MSI Growth Rates



NOTE: The left panels show month-to-month percentage change, at an annual rate, in each MSI aggregate; changes in excess of ±30 percent are excluded. Excluded dates and values are M1, 1986:12 (32.4 percent), 2001:09 (44.8 percent), 2008:09 (32.9 percent), 2008:12 (50.9 percent); M2M, 2001:09 (37.2 percent), 2008:12 (32.5 percent); M2, 2001:09 (32.7 percent), 2008:12 (31.1 percent); MZM, 2001:09 (38.9 percent), 2008:12 (32.8 percent); ALL, 2001:09 (34.7 percent), 2008:12 (31.7 percent). Right panels show year-to-year percentage change, monthly (no truncated points are omitted). Absolute decreases in MSI-MZM and MSI-ALL during 2010 were due to sharp decreases in institutional-type MMMFs. Intervals between National Bureau of Economic Research (NBER) business cycle peaks and troughs are shaded.

Figure 9

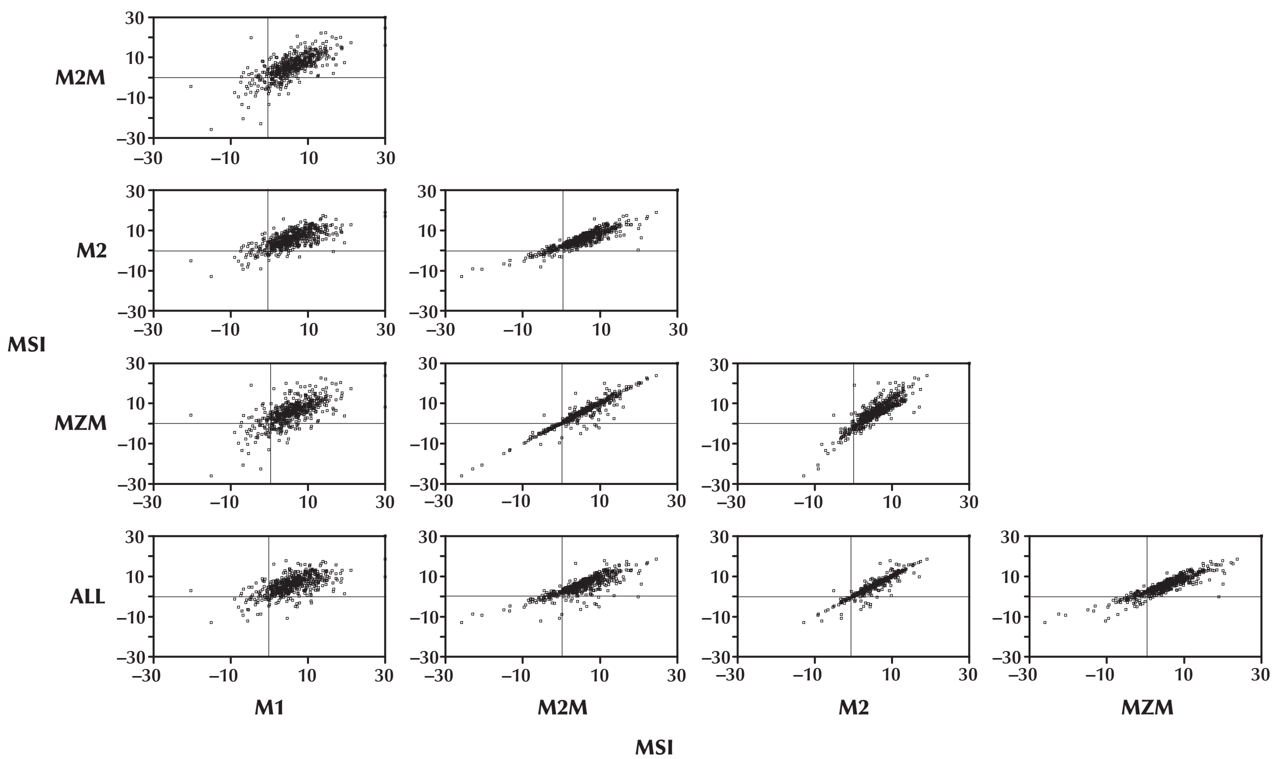
Growth of the MSI Aggregates (selected decades)



NOTE: Each panel compares the path of (the log of) the specified MSI aggregate during a selected decade-long period. Intervals between NBER business cycle peaks and troughs are shaded. For A, log levels, January 1967 = 10. For B, log levels, January 1978 = 1.0. For C, log levels, January 1989 = 1.0. For D, log levels, March 2001 = 1.0.

Figure 10

MSI Indexes (January 1967–April 2011)



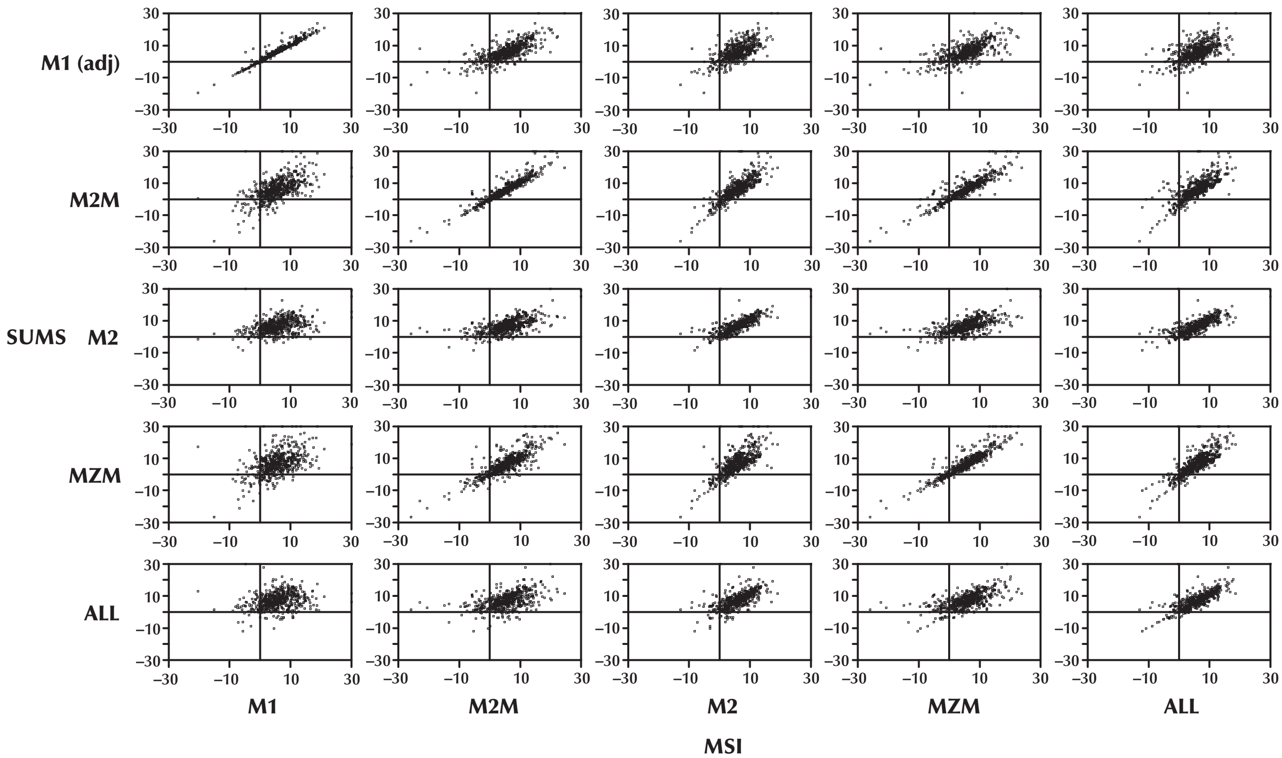
NOTE: Each panel compares the monthly growth of the two specified MSI. The MSI aggregates are adjusted for the effects of retail deposit sweep programs by removing such deposits from savings deposits and including them in checkable deposits. Monthly growth rates in excess of 30 percent are excluded from the chart; see the footnote to Figure 8 for a list of the excluded points.

top panel of Table 3. We note that M3 is included in Table 3 but not in the associated figures. Among the MSI, the lowest correlations involve MSI-M1 (first column). Aside from MSI-M1, all of the MSI are mutually highly correlated. The correlation between MSI-M2M and MSI-MZM is 0.95 and that between MSI-M2 and MSI-ALL is 0.94. The middle panel of Table 3 shows the correlations between the sum aggregates, which (except for M2M and MZM) are lower than the correlations among the MSI. Figure 11 shows a scatterplot matrix comparing the month-to-month growth rates of the MSI and the corresponding summation aggregates; corresponding correlations are shown in the bottom panel of Table 3.When comparing SUM aggregates with the MSI, a general finding is that the SUM aggregates are all most highly correlated with the MSI constructed over the same components.[33] The correlation is highest (0.98) between SUM-M1 (adjusted for retail deposit sweeps) and MSI-M1. For M2M, M2, MZM, and ALL, the correlations between the SUM aggregate and the comparable MSI are between 0.78 and 0.80. For M3, it is 0.69. SUM-M1 (adjusted for retail sweeps) displays modest positive correlation with all of the other MSI as well (0.64 to 0.75). Other SUM aggregates show modest positive correlation with at least some MSI, although correlations involving MSI-M1 are all low. For example, correlations between SUM-M2

[33] Except for M3, the MSI are also most highly correlated with the SUM aggregate constructed over the same components.

Figure 11

MSI Indexes Versus Sum Money Aggregates (January 1967–April 2011)



NOTE: Each panel compares the monthly growth of the specified MSI and the corresponding simple-sum monetary aggregate. The sum aggregate M1 (adj) and all MSI are adjusted for the effects of retail deposit sweep programs as described in the text. Excluded points for the MSI are listed in the notes of Figure 8. Excluded points for the Sum aggregates are for M1 adjusted for retail sweeps, 1986:12 (32.7), 2001:09 (43.2), 2008:09 (32.5), 2008:12 (50.1); for M2M, 1982:12 (44.4), 1983:01 (123.8), 1983:02 (79.1), 1983:03 (35.2), 2001:09 (32.7), 2008:12 (30.8); for M2, 1983:01 (33.40); for MZM, 1980:07 (31.4), 1981:04 (30.1), 1982:12 (41.0), 1983:01 (116.9), 1983:02 (75.7), 1983:03 (32.5), 2001:09 (37.3), 2008:02 (31.4), 2008:12 (31.6); for ALL, 1983:01 (31.7), 2001:09 (30.2), 2008:12 (30.1). Values in parentheses indicate percent.

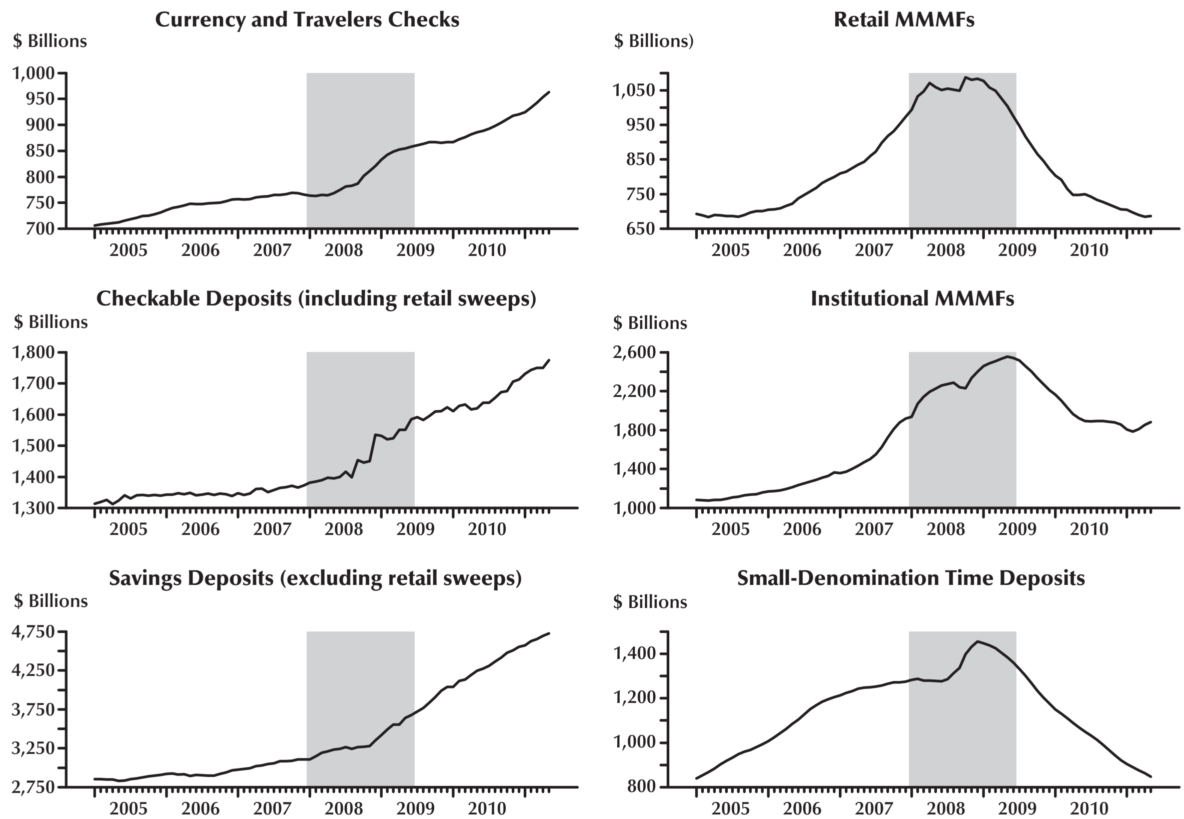
and MSI other than MSI-M1 range from 0.62 to 0.75. It is interesting to note all of the MSI, except MSI-M3, are less highly correlated with MSI-M1 than they are with the SUM aggregate constructed over the same components (for MSI-M3, the correlations are the same). Thus, the choice between M1 and a broader level of aggregation clearly matters more than the method of aggregation.

A Case Study of MSI Components During the 2007-2009 Financial Crisis

Economic theory suggests that the MSI provide a superior measure of the flow of monetary services that households and firms derive from holding monetary assets. In this section, we briefly survey the behavior of component quantities of the MSI and their user costs during the financial crisis; more complete modeling is a topic for future research. The construction of the MSI provides data beyond quantity indexes, including measures of the user cost of monetary services (essentially, the price of owning immediately available funds plus the cost of positioning an asset portfolio so that such funds are readily available when needed) that are interesting to examine because the recent financial crisis was largely a "liquidity" crisis in

Figure 12

Major Components of MSI



NOTE: Each panel shows the nominal quantity (in billions of current dollars) of a group of monetary assets included in the MSI. Checkable deposits includes, and savings deposits excludes, the amount of deposits in retail deposit sweep programs.

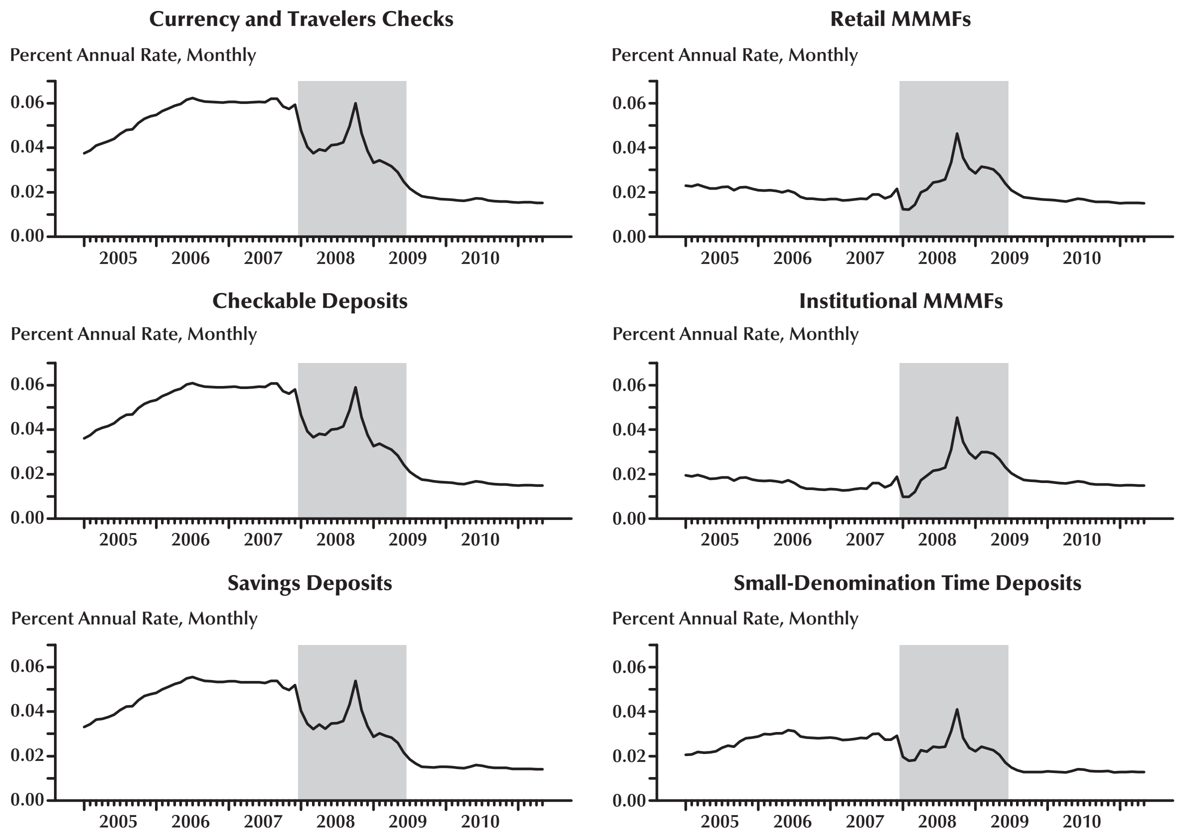
which the functioning of short-term money and financial markets was impaired.

Figure 12 shows changes in the component quantities of the MSI. For checkable deposits (demand deposits and OCDs), savings, and small-denomination time deposits, we have summed the amounts at commercial banks and thrift institutions, calculating the user costs by dividing the corresponding summed expenditures by the sum of the quantities. There is a clear break in all series circa September 2008 (when Lehman Brothers filed for bankruptcy, the federal government rescued AIG, and the shadow banking system largely shut down).[34] With respect to individual components during late 2008, the sharp increase in checkable deposits reflects, in part, extension by the FDIC in October 2008 of deposit insurance without limit to non-interest-bearing checkable deposits. Increases in savings and small-denomination time deposits likely reflect flight to quality since these deposits are FDIC insured. Small dimples in MMMF shares likely reflect shaken investor confidence before the federal government's October 2008 de facto extension of deposit insurance to shares in these funds. During 2009 and 2010, steep runoffs in the components except currency, checkable deposits, and savings deposits are

[34] Anderson and Gascon (2009) examine the 2008 sudden stop in the asset-backed commercial paper market, the heart of the U.S. "shadow" banking system.

Figure 13

Real User Cost of Monetary Services by Component



NOTE: Each panel shows the real user cost of the monetary services obtained through the specified group of assets. Within each group, the user costs are simple arithmetic averages of the user costs of the components within that group.

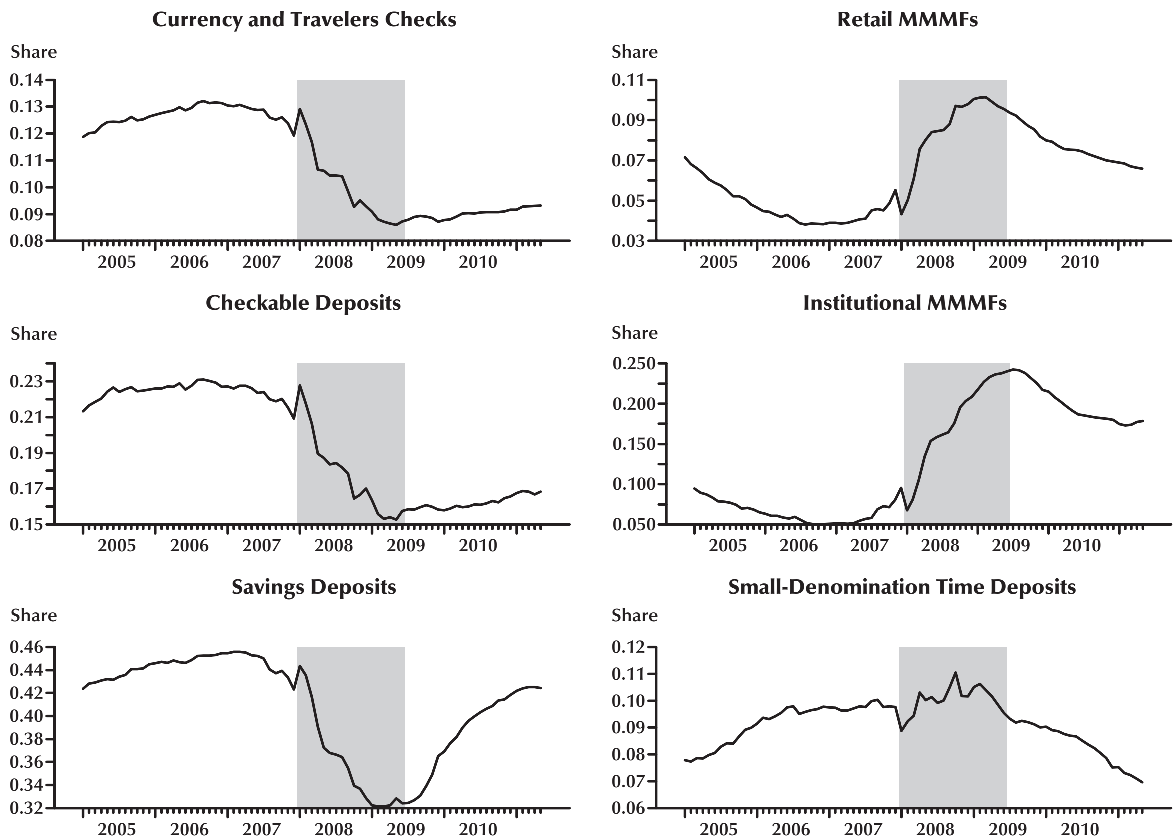
apparent; the trough in economic activity occurred during the second quarter of 2009 and the stock market low point was in March 2009.

Whereas economic theory suggests that all components of an MSI furnish monetary services (either as medium of exchange or as an asset readily convertible to medium of exchange), the user costs (prices) of monetary assets differ because they earn different rates of return. Figure 13 shows the real user costs. Three major shifts are apparent. First, the user costs decreased sharply in late 2007–early 2008, in part due to FOMC policy actions. Second, the user costs increased almost steadily during 2008 before jumping sharply during September when short-term financial markets shut down. Third, user costs decreased sharply and then steadily during 2009, again in response to FOMC policy action. During the financial crisis, the user cost of monetary services accurately tracked the scarcity of liquidity due to impaired short-term money markets. Indeed, increases in the price of monetary services (i.e., the price of liquidity) reinforce Federal Reserve Chairman Ben Bernanke's argument that immediate relief of financial market stress ("credit easing") was essential.[35]

[35] Anderson and Gascon (2011) survey many such market assistance programs and review evidence regarding their role in preventing a slide into economic depression.

Figure 14

Expenditure Shares of Monetary Services



NOTE: Each panel shows expenditures on a set of components from MSI-ALL as a share of total expenditure on MSI-ALL. Expenditures incurred to obtain the monetary services furnished by each group of assets equal the sum (across the assets in the group) of the products of each asset's real user cost times the asset's nominal quantity (or equivalently, the product of each asset's nominal user cost times the real quantity of each asset).

Figure 14 combines information from Figures 12 and 13 to examine fluctuations in expenditures on money services during the financial crisis. Note that the behavior of component expenditure *shares* depends on the price elasticity of demand—a component's share might increase, decrease, or be unchanged when its price increases. In the left column of the figure, the shares of expenditure (on monetary services) corresponding to the three most-liquid assets decreased throughout 2008. In contrast, the expenditure shares for MMMFs increased throughout 2008. The expenditure share for small-denomination time deposits increased slightly during 2008, followed by a sharp decrease during 2009. Deeper understanding of these expenditure share patterns is a topic for future research.[36]

SUMMARY AND CONCLUSIONS

The revised MSI and the underlying data used to construct them as presented in this article are

[36] A standard way to model the expenditure share system would be to estimate a demand systems model for the monetary assets. Jones et al. (2008), for example, estimate a Fourier flexible functional form model for the components of MSI-M2.

valuable resources to empirical economists interested in the role that money plays in the economy. This revision to the MSI includes a number of changes that should improve the quality of the underlying data. We now adjust the savings and checkable deposit components for the effects of retail sweep programs and have improved the measures of savings and small-denomination time deposit rates in the Regulation Q era. We also introduce a new benchmark rate, defined as the largest rate in a set of rates that includes the own rates of the components of the broadest index and yields on selected short-term money market rates (the upper envelope) plus a modest liquidity premium. Previously the benchmark rate was defined by including the Baa bond yield in the upper envelope and not adding a liquidity premium. We believe that this new benchmark rate improves on the previous practice since recent literature has criticized using long-term bond yields to measure the benchmark rate (e.g., Stracca, 2004). Nevertheless, the benchmark rate remains one of the most problematic aspects of the Divisia approach in terms of being as much a conceptual issue as an empirical one.

Several issues raised in this paper relate to the components of M1. A major problem with the official M1 monetary aggregate is, of course, retail sweeping of transaction deposits. Although not addressed in this paper, commercial sweeping of demand deposits further complicates interpretation of the M1 aggregate (see Jones, Dutkowsky, and Elger, 2005, and Dutkowsky, Cynamon, and Jones, 2006). Another problem concerns implicit remuneration of demand deposits. Within M1, only OCDs are explicitly remunerated. Before the paper by Anderson, Jones, and Nesmith (1997c), the builders of Divisia monetary aggregates proceeded as if demand deposits held by businesses were remunerated at a rate equal to the short-term commercial paper rate adjusted for statutory reserve requirements; the data used to distinguish between business and household demand deposits were discontinued in 1990. Our decision to not include implicit returns on demand deposits is based on this lack of data. We are concerned, in particular, about whether MSI-M1 provides much useful information beyond what is provided by a sweep-adjusted version of its simple sum counterpart (e.g., Figure 11, row 1). Moreover, the main advantage of the Divisia approach relates to the construction of broader aggregates. As Lucas (2000, pp. 270-71) has argued, "I share the widely held opinion that M1 is too narrow an aggregate for this period [the 1990s], and I think that the Divisia approach offers much the best prospects for resolving the difficulty."

Finally, a number of analysts have argued that the Federal Reserve's 2006 decision to discontinue publication of its M3 monetary aggregate significantly constrained subsequent monetary analyses. We found that the growth rates of MSI-M2 and MSI-ALL are both highly correlated with growth rates of MSI-M3 when it can be computed providing some reassurance with respect to the discontinuance of M3. On the other hand, growth rates of MSI-M2 and MSI-ALL diverged much more than usual in 2010, suggesting that MSI-M3 might have contained some additional information in recent years.

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TECHNICAL APPENDIX
More on Index Number Theory

In index number theory, a *quantity index* is a function of the quantities and prices of a set of goods in two periods. Diewert (1976a) defined the concepts of exact and superlative indexes, which Barnett (1980) applied to monetary data.[37] In this context, the real stocks of a set of monetary assets act as quantities paired with their corresponding user cost prices. With this in mind, let $Q(\pi_0,\pi_1,m_0,m_1)$ denote a *monetary* quantity index, where π_0,m_0 and π_1,m_1 represent data for a set of monetary assets in two periods. Following Diewert (1993, p. 198), the quantity index is *exact* for the utility function $V(m)$ if

$$V(m_1) = V(m_0)Q(\pi_0,\pi_1,m_0,m_1)$$

for every π_0,π_1,m_0,m_1 such that m_i maximizes $V(m)$ subject to $\pi_i \cdot m \le \pi_i \cdot m_i$ for $i = 0,1$. Thus, if the quantity index is greater (less) than 1, it indicates that the utility function is increasing (decreasing) from m_0 to m_1. A quantity index is *superlative* if it is exact for a linearly homogeneous utility function, which can provide a second-order approximation to an arbitrary linearly homogeneous utility function (see Diewert, 1993, p. 204).

Diewert (1976a) also defined exact and superlative price indexes. A quantity index, Q, and a price index, P, satisfy weak factor reversal (Diewert, 1993, p. 198) if

$$Q(\pi_0,\pi_1,m_0,m_1)P(\pi_0,\pi_1,m_0,m_1) = \frac{\pi_1 \cdot m_1}{\pi_0 \cdot m_0}.$$

Following Diewert (1993, p. 206), a pair of indexes is superlative if either the quantity index is superlative or the price index is superlative and they satisfy weak factor reversal.

Chain-weighting can be used to facilitate comparisons over more than two periods. A *chain-weighted quantity index*, V_t, is a time series constructed from a quantity index number formula as follows:

$$V_t = V_{t-1}Q(\pi_{t-1},\pi_t,m_{t-1},m_t).$$

The MSI are chain-weighted monetary quantity indexes, which are based on the superlative Törnqvist-Theil quantity index number formula. An alternative would be to use the Fisher ideal quantity index, which is also superlative.[38] Farr and Johnson (1985), for example, used the Fisher ideal index to construct their series. In principle, other superlative indexes could also be considered. Diewert (1976a) proved that the quadratic mean of order r quantity and price indexes is superlative for all r. The Törnqvist-Theil and Fisher ideal indexes are special cases ($r = 0$ and $r = 2$, respectively). Using two datasets, Hill (2006, p. 38) found that these superlative indexes can differ significantly from one another. His interpretation is as follows:

> The problem is that, as the parameter r increases in absolute value, the superlative price (quantity) index formula becomes increasingly sensitive to outliers in the price-relatives (quantity-relatives) distribution. Admittedly, the only three superlative indexes that have been seriously advocated in the index number literature all lie in the range $0 \le r \le 2$, where the superlative formula is relatively unaffected by outliers. Over this range, all the superlative indexes do approximate each other closely. Conceptually, however, the economic approach provides no reason for restricting the range thus.

Thus, the choice of which index number formula to use seems to be more consequential than previously thought.

[37] The reader is referred to Diewert (1993) for a detailed survey of index number theory.

[38] Both the Törnqvist-Theil and Fisher ideal quantity indexes have also been shown to have attractive properties even if the utility function is not linearly homogeneous (see Diewert, 1976a,b, and Diewert, 1993, pp. 211-13). Diewert (1976a, pp. 137-38) offers several reasons why the Fisher ideal index is preferable to other superlative indexes; among these is its consistency with revealed preference theory.

A Survey of Announcement Effects on Foreign Exchange Volatility and Jumps

Christopher J. Neely

This article reviews, evaluates, and links research that studies foreign exchange volatility reaction to macro announcements. Scheduled and unscheduled news typically raises volatility for about an hour and often causes price discontinuities or jumps. News contributes substantially to volatility but other factors contribute even more to periodic volatility. The same types of news that affect returns—payrolls, trade balance, and interest rate shocks—are also the most likely to affect volatility, and U.S. news tends to produce more volatility than foreign news. Recent research has linked news to volatility through the former's effect on order flow. Empirical research has confirmed the predictions of microstructure theory on how volatility might depend on a number of factors: the precision of the information in the news, the state of the business cycle, and the heterogeneity of traders' beliefs. (JEL F31, E01, E44)

Federal Reserve Bank of St. Louis *Review*, September/October 2011, *93*(5), pp. 361-407.

Researchers have long sought to understand how announcements of various sorts affect foreign exchange *volatility*, which is the magnitude of changes in foreign exchange rates. Unfortunately, such studies are frequently disconnected from each other, making it difficult for casual observers to see the big picture. To remedy this situation, this paper surveys and draws together the literature on announcements and foreign exchange volatility.[1]

The literature on announcements and foreign exchange volatility is part of a larger literature that seeks to characterize patterns in conditional variance or conditional standard deviations (SDs). People and firms do not like volatile asset prices because they are risk averse; loss of wealth puts their desired consumption at risk. Similarly, traders must quantify the volatility of their positions because excessive losses put their jobs at risk. Understanding and estimating asset price volatility is therefore important for asset pricing, portfolio allocation, and risk management.

Asset price volatility can change for a variety of reasons: the opening or closing of markets, a changing rate of news arrival, or a change in the rate of how agents act on information. Together, these factors produce three prominent characteristics in foreign exchange volatility: (i) It tends to be autocorrelated; (ii) it is periodic, displaying intraday and intraweek patterns; and (iii) it includes discontinuities (jumps) in prices.

Characterizing asset price volatility is an important goal for financial economists. Scheduled macroeconomic announcements are useful natural experiments through which to study how the release of public information affects prices and volatility. Because survey expectations permit researchers to measure the surprise component

[1] A companion article, Neely and Dey (2010), surveys the related literature on the response of foreign exchange returns to announcements.

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of an announcement, researchers can distinguish the reaction of volatility to the magnitude of surprises from the reaction of volatility to the existence of the announcement itself.

Prior to the formal study of announcement effects on volatility, researchers found that volatility is autocorrelated and displays intraday and intraweek patterns. In addition, many regularly scheduled announcements—especially those that affected returns—also influenced volatility. Researchers sought to distinguish patterns caused by market opening/closing from those caused by regular macro announcements. Although Ederington and Lee (1993) argue that announcements account for most intraday and intraweek volatility patterns, Andersen and Bollerslev (1998) demur; they stress that it is important to jointly model the contributions of announcements, other intraday patterns, and the persistent component of volatility.

The study of announcements and volatility also has direct implications for policy. For example, some policy analysts have proposed taxing foreign exchange transactions to reduce allegedly meaningless churning that creates "excess" volatility. Melvin and Yin (2000), however, establish a strong link between news arrival and volatility, which argues against proposals to reduce trading volume through regulation.

Much of the literature on volatility patterns and news is only loosely linked to microstructure theory; it seeks mainly to characterize which announcements are important influences on volatility and how long the effects last. At times, however, microstructure theory has influenced the study of announcement effects on volatility, volume, and spreads. For example, microstructure theory motivates the study of how market conditions—heterogeneity of interpretation or the presence of conflicting information or the state of the business cycle, or the quality of information—influence reactions to announcements (Baillie and Bollerslev, 1991, and Laakkonen and Lanne, 2009). More recently, researchers have considered the relative importance of public and private information releases in creating price volatility through order flow (Cai et al., 2001; Evans, 2002; Evans and Lyons, 2005).

This survey considers the impact of announcements on price discontinuities (jumps) because jumps are defined by their magnitude and have implications for volatility forecasting. Specifically, Neely (1999) and Andersen, Bollerslev, and Diebold (2007) show that removing jumps from current and lagged volatility estimates improves the accuracy of volatility forecasts.

The next section begins with a discussion of the methodological considerations involved in studying the effect of announcements on volatility. This is followed by a review of the major areas of research on the effect of announcements on foreign exchange market volatility. The final section includes a discussion of the results and conclusions.[2]

METHODS OF STUDYING ANNOUNCEMENT EFFECTS ON FOREIGN EXCHANGE VOLATILITY

Methodology

Two methodological questions arise in the study of the effects of announcements on foreign exchange volatility: How should volatility be measured, and what information about announcements influences volatility? Researchers have used three measures of volatility to study announcement effects: implied volatility, which is an estimate of future volatility derived from option prices; high-frequency squared returns, a nonparametric method that Andersen and Bollerslev (1998) later formalized as realized volatility; and volatility estimated parametrically by some variant of generalized autoregressive conditionally heteroskedastic (GARCH) models (Engle, 1982, and Bollerslev, 1986).[3]

Volatility measures respond differently to macro announcements because they approach

[2] Neely and Dey (2010) describe the most commonly studied U.S. announcements.

[3] Neely (2005) discusses the measurement and uses of implied volatility estimated from options prices. Engle (1982) developed the autoregressive conditionally heteroskedastic (ARCH) model that Bollerslev (1986) extended to the GARCH formulation. GARCH models usefully account for the time-varying volatility and fat-tailed distributions of daily and intraday financial returns.

volatility in different ways. *Implied* volatility, for instance, approximates average volatility until the expiry of the option, which could be in weeks or months. Therefore, it is strongly forward looking and often insensitive to short-lived volatility effects from macro announcements. Likewise, GARCH models fit to daily data predict daily volatility through essentially autoregressive processes, but such models cannot estimate intraday effects. In contrast to implied volatility or daily GARCH estimates, high-frequency data—which can be used with parametric models such as GARCH—are well suited to measuring short-lived, intraday effects.

The second issue is what type of information about announcements influences volatility. A scheduled announcement itself—regardless of content—could be expected to change volatility either before or after the announcement. In addition, surprising information in the announcement might influence volatility by precipitating additional trading from revised expectations. In practice, researchers have used both announcement indicators and surprises, sometimes finding different effects.

An announcement is "surprising" to the extent that it deviates from market expectations. To construct announcement surprises, researchers generally use the median response from the Money Market Services (MMS) survey to estimate the expected announcement. Each Friday, MMS surveys 40 (formerly 30) money managers on their expectations of forthcoming economic releases.[4] Cornell (1982) and Engel and Frankel (1984) first used these survey data in the literature on announcement effects in the foreign exchange market, though other researchers (e.g., Grossman, 1981) had used them in other contexts. Grossman (1981), Engel and Frankel (1984), Pearce and Roley (1985), and McQueen and Roley (1993) show that the MMS survey data estimate news announcements in an approximately unbiased and informationally efficient fashion, outperforming time-series models.[5]

To compare coefficients on announcement surprise series with different magnitudes, researchers have typically followed Balduzzi, Elton, and Green (2001) in standardizing surprises by subtracting the MMS expectation from the release and dividing those differences by the SD of the series of differences. For example, the standardized surprise for announcement *j* is as follows:

$$(1) \quad S_t^j = \frac{R_t^j - E_t^j}{\hat{\sigma}_j},$$

where R_t^j is the realization of announcement j on day t, E_t^j is the MMS market expectation, and $\hat{\sigma}_t$ is the estimated SD of the series of the differences.[6] Thus, announcement surprises are close to mean zero and have a unit SD.

THE LITERATURE ON ANNOUNCEMENTS AND FOREIGN EXCHANGE VOLATILITY

Early Study of Volatility Patterns

The earliest studies of announcement effects on the foreign exchange market considered only the reaction of prices/returns, but researchers added focus on volatility in the 1990s. Early studies of volatility patterns by Engle, Ito, and Lin (1990) and Harvey and Huang (1991) motivated this work, although the latter paper did not explicitly incorporate macro announcements.

Harvey and Huang (1991) discover an intraday U-shaped volatility pattern in hourly foreign exchange returns as well as intraweek effects. Volatility is higher on Thursday and Friday but

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[4] The number of survey participants and the dates of the survey have changed over time. Hakkio and Pearce (1985) report that MMS surveyed about 60 money market participants during that era and that they conducted the surveys on both Tuesdays and Thursdays before February 8, 1980, and on Tuesdays after that date.

[5] Although the MMS survey expectations exhibit fairly good properties compared with alternatives, they still surely measure market expectations with some error, both because they are at least a couple days old and because they reflect the views of a small group of money managers. More subtly, any macroeconomic release will surely contain some error about the true state of the economy because it is estimated with finite resources and limited information. Therefore, the macroeconomic surprise will be estimated with error and this error will generally attenuate the estimated market response toward zero. Rigobon and Sack (2008) discuss two methods to compensate for this error. Bartolini, Goldberg, and Sacarny (2008) discuss the application of this methodology.

[6] In a personal communication, Mike McCracken raises the interesting question of whether it would be better to normalize with the conditional SD.

volatility on Monday is no different from volatility on Tuesday. The authors speculate that important news announcements at the end of the week raise volatility on Thursday and Friday. Finally, volatility is highest during the traded currency's own domestic business hours, particularly so for non-USD (U.S. dollar) cross rates. For example, USD volatilities peak during U.S. trading hours, implying the potential importance of U.S. macroeconomic announcements (Ito and Roley, 1987).

Engle, Ito, and Lin (1990) extend this research in intraday volatility patterns by introducing the concepts of *heat waves* and *meteor showers* in the foreign exchange market. Heat waves refer to the idea that volatility is geographically determined—that is, a heat wave might raise volatility in New York on Monday and Tuesday but not in London on Tuesday morning. Heat waves might occur if most or all important news that affects volatility occurs during a particular country's business day and there is little price discovery when that country's markets are closed. In contrast, meteor showers refer to the tendency of volatility to spill over from market to market, from Asian to European to North American markets, for example. Therefore, meteor showers imply volatility clusters in time, not by geography. Using a GARCH model with intraday data, Engle, Ito, and Lin (1990) find that the meteor shower hypothesis better characterizes foreign exchange volatility engendered by balance of trade announcements.[7] Baillie and Bollerslev (1991) confirm the meteor shower effect but also find some evidence of heat wave behavior.

Motivated by the microstructure theory of Epps and Epps (1976) and Tauchen and Pitts (1983), Hogan and Melvin (1994) follow up on the meteor shower/heat wave literature by exploring the role of heterogeneous expectations in volatility persistence across markets. Using the SD of MMS responses to measure heterogeneity of market expectations in a four-observations-per-day GARCH model, Hogan and Melvin (1994) find support for the idea that heterogeneous expectations *do* increase volatility persistence in the wake of a U.S. trade balance announcement.[8] In retrospect, it seems unsurprising that meteor showers should predominate over heat waves in a world of global trading and a high degree of autocorrelated common shocks across countries: News tends to cluster in time and will surely affect volatility across the globe.

Early Research on Announcements and Volatility

Harvey and Huang (1991) and the meteor shower/heat wave literature found intraday and intraweek patterns that indicated that macro announcements were potentially important sources of volatility. Later studies extended this research by directly studying the effect of announcements on various measures of foreign exchange volatility.

In the late 1980s and early 1990s, U.S. trade deficit news was considered very important, especially for the USD/JPY (Japanese yen) exchange rate. Two of the earliest papers examine volatility responses to these releases. Madura and Tucker (1992) analyze the effect of trade balance surprises on the change in average implied SDs (volatilities) of currency options from the day before the announcement to the day of the announcement. They argue that studying implied volatility permits researchers to observe how announcements change the market's (long-run) ex ante volatility forecast. Although unexpected news—good or bad—increases implied volatilities, the announcement itself tends to reduce them. This probably reflects the fact that implied volatilities look forward over several months. While announcements generally increase volatility over the very short term, resolving the uncertainty associated with the announcement should reduce expected volatility over longer horizons.

Using a bivariate GARCH model to study spot and futures market responses to U.S. trade deficit announcements, Sultan (1994) finds two types of asymmetry in daily volatility responses: The USD/JPY is much more responsive to trade

[7] The appendix describes the key features of the papers studying announcement effects on volatility.

[8] Curiously, Hogan and Melvin (1994) find that news has no impact on conditional volatility. This is almost certainly due to a misspecification; the authors specify conditional volatility as a function of signed news surprises rather than absolute news surprises.

deficit news than other exchange rates, and larger-than-expected U.S. trade deficits provoke much stronger volatility responses than smaller-than-expected ones, presumably because larger trade deficits are much more likely to provoke a policy response than smaller deficits.

In contrast to the work with implied volatility and daily GARCH modeling, Ederington and Lee (1993, 1994) investigate how U.S. macroeconomic release indicators affect very short-run volatility: absolute 5-minute USD/DEM (German deutsche mark) and USD/JPY returns, respectively. The merchandise trade deficit, employment report, producer price index (PPI), durable goods orders, gross national product (GNP), and retail sales all affect USD/DEM volatility significantly.[9] Volatility is not particularly high at the opening of the market (8:20 a.m. ET) but increases 10 minutes later at 8:30 a.m., which is the time of many major announcements. It remains very high for 15 minutes and higher than normal for several hours following a news release. After controlling for announcement effects, the authors find that average volatility is flat over both the trading day and week—that is, news "mainly" explains both intraday and weekly patterns. Using 10-second data, Ederington and Lee (1995) observe high USD/DEM futures volatility immediately preceding a news announcement but find no evidence of information leakage. Volatility might anticipate news surprises.

Decomposing Announcements and Periodic Volatility Patterns

The very early literature on announcements and volatility noted the periodicity in volatility and speculated that announcements might be responsible. The work of Ederington and Lee (1993, 1994, and 1995) illustrated the importance of announcements for volatility and considered whether there was any residual, unexplained periodicity: "We find these [macro] announcements are responsible for most of the observed time-of-day and day-of-the-week volatility patterns in these [foreign exchange] markets" (Ederington and Lee, 1993, p. 1161).

Because announcements and periodicity are correlated, however, one must jointly model them to consistently estimate and compare their impact (Payne, 1996, and Andersen and Bollerslev, 1998). In particular, Andersen and Bollerslev (1998) use 5-minute USD/DEM currency returns to integrate prior research on daily volatility persistence, intraday and intraweek periodicity, and announcement effects. They affirm the importance of macro releases as addressed by Ederington and Lee (1993), but argue that these are secondary to the intraday pattern; periodic patterns and autoregressive volatility forecasts explain more of intraday and daily volatility than do announcements.

Presaging the literature on the effect of announcements on order flow, Andersen and Bollerslev (1998) conjecture that the intraday volatility pattern alters daily trading patterns. Further, they find that—after accounting for the intraday volatility pattern—including ARCH terms significantly improves forecasting power, even in a high-frequency volatility process.[10] Real U.S. announcements—employment, gross domestic product (GDP), trade balance, and durable goods orders—are the most influential U.S. announcements in explaining volatility movements, while monetary policy news is most significant among German announcements. This finding is consistent with the conventional wisdom that the Bundesbank was relatively more concerned with monetary measures than the Federal Reserve.

The debate on the relative importance of pure periodicity versus announcement effects continued after publication of the paper by Andersen and Bollerslev (1998). To compare periodicity to announcement effects on foreign exchange volatility, Han, Kling, and Sell (1999) and Ederington and Lee (2001) both examine USD futures data, finding similar results but interpreting them differently. Using high-frequency futures data for four currencies from 1990 to 1997, Han, Kling, and Sell (1999) show that the DEM and JPY exhibit strong day-of-the-week volatility effects, even after controlling for indicators of 18 U.S. announcements. These authors speculate that differences in their testing procedures—testing

[9] Leng (1996) notes that major announcements have longer-lived effects on volatility than minor announcements.

[10] Andersen and Bollerslev (1998) argue that the intraday volatility pattern obscures ARCH effects in intraday data.

by interval, rather than over pooled intervals—might account for the disparity in their conclusions with those of Ederington and Lee (1993). Ederington and Lee (2001) compare the power of seasonal effects, macro announcement indicators, and past volatility to predict volatility in 10-minute futures data on the DEM/USD from July 1989 through May 1993. Confirming their 1993 research but disputing the inference of Andersen and Bollerslev (1998), Ederington and Lee (2001) argue that macro announcements create most of the time-of-day and day-of-week effects and greatly reduce persistence in ARCH models. Unscheduled announcements create volatility that persists longer than that of scheduled announcements.

The appearance of contradictory results is at least partly due to a difference in emphasis: Ederington and Lee (2001) argue that announcements are more important than day-of-the-week effects, but Han, Kling, and Sell (1999) take the null hypothesis to be no day-of-the-week effects after controlling for announcements. The use of futures data by both studies, however, is likely to bias the results in favor of the importance of announcements, as the futures markets are open for U.S. announcements but not for important periodic shifts in volatility during non-U.S. business hours.

How can we resolve the disparate conclusions of Andersen and Bollerslev (1998) and Ederington and Lee (2001) about the relative importance of announcement effects and other periodic factors? To illustrate the issues involved in disentangling announcement and other periodic effects, one can regress absolute hourly foreign exchange returns—24 hours a day, 5 days a week—on announcement variables and periodic components. The following equation describes such a regression for hourly returns:

(2)

$$\begin{aligned} \left|r_t\right| &= \alpha + \beta_{1,US} Dum_{USann,t} + \beta_{1,for} Dum_{forann,t} \\ &+ \sum_{j=1}^{N} \beta_{2,j} \left|s_{j,t}\right| + \sum_{q=1}^{4} \left(\beta_{3,q} \cos\left(\frac{q2\pi t}{24}\right) + \beta_{4,q} \sin\left(\frac{q2\pi t}{24}\right) \right) \\ &+ \sum_{\iota=1}^{5} \beta_{5,\iota} \left|r_{t-\iota}\right| + \beta_6 \frac{\widehat{\sigma_{d(t)}}}{\sqrt{24}} + \sum_{h=19}^{23} \beta_{7,h} Dum_{FRI_h,t} + \varepsilon_t, \end{aligned}$$

where r_t is the annualized log return from period t to t+1; $Dum_{USann,t}$ and $Dum_{forann,t}$ are dummy variables that take the value 1 if there is any U.S. or foreign announcement, respectively, during t to t+1, and 0 otherwise; $s_{j,t}$ is the standardized surprise of announcement j at period t;

$$\cos\left(\frac{q2\pi t}{24}\right) \text{ and } \sin\left(\frac{q2\pi t}{24}\right)$$

are trigonometric functions that allow parsimonious estimation of an intraday periodic component; and

$$\widehat{\sigma_{d(t)}}$$

is the square root of the 1-day-ahead annualized GARCH(1,1) daily volatility forecast for day $d(t)$.[11] Finally, $Dum_{FRI_h,t}$ takes the value 1 if period t coincides with hour h of a Friday, and 0 otherwise. The treatment of periodicity in equation (2) differs from that of either Han, Kling, and Sell (1999) or Ederington and Lee (2001), who both used less-parsimonious combinations of indicator variables for times of the day. Equation (2) is closer in spirit to the work of Andersen and Bollerslev (1998).

I estimate equation (2) by ordinary least squares on 1-hour log changes in the USD/EUR (euro) exchange rate over the period November 5, 2001, to March 12, 2010, after first removing weekends and the following holidays from the sample: New Year's Day (December 31–January 2), Good Friday, Easter Monday, Memorial Day, Fourth of July (July 3 or 5 when the Fourth falls on a Saturday or Sunday, respectively), Labor Day, Thanksgiving (and the Friday after), and Christmas (December 24-26).

Table 1 shows the relative explanatory power of the various components of equation (2) for absolute returns. The full regression has a substantial R^2 of 0.2211, with the greatest explanatory power coming from the intraday periodicity with a partial R^2 of 0.0514, and the GARCH daily volatility forecast (0.0429). The announcement dummies provide a partial R^2 of 0.0020 and the

[11] Equation (2) could be altered to take into account a host of effects, including asymmetry or business cycle dependence, for example.

Table 1

R^2 and Partial R^2s

Independent variable(s)	R^2 or partial R^2s
Full regression	0.2211
Seasonal effect	0.0514
GARCH(1,1) volatility forecast	0.0429
Absolute announcement surprises	0.0199
Lags of absolute returns	0.0156
Friday night dummy variables	0.0103
Announcement dummies	0.0020

NOTE: The table displays the R^2 and partial R^2s from regression (2) and various combinations of its regressors: the announcement dummies, $\beta_{1,US}Dum_{USann,t}$ and $\beta_{1,for}Dum_{forann,t}$; the absolute announcement surprises, $\sum_{j=1}^{N}\beta_{2,j}\left|s_{j,t}\right|$; the periodic component, $\sum_{q=1}^{4}\left(\beta_{3,q}\cos\left(\frac{q2\pi t}{24}\right)+\beta_{4,q}\sin\left(\frac{q2\pi t}{24}\right)\right)$; five lags of absolute returns, $\sum_{\iota=1}^{5}\beta_{5,\iota}\left|r_{t-\iota}\right|$; the GARCH (1,1) daily volatility forecast, $\beta_6\frac{\widehat{\sigma_{d(t)}}}{\sqrt{24}}$; and the Friday night indicators, $\sum_{h=19}^{23}\beta_{7,h}Dum_{FRI_h,t}$.

$$\left|r_t\right| = \alpha+\beta_{1,US}Dum_{USann,t}+\beta_{1,for}Dum_{forann,t}+\sum_{j=1}^{N}\beta_{2,j}\left|s_{j,t}\right|+\sum_{q=1}^{4}\left(\beta_{3,q}\cos\left(\frac{q2\pi t}{24}\right)+\beta_{4,q}\sin\left(\frac{q2\pi t}{24}\right)\right)$$
$$+\sum_{\iota=1}^{5}\beta_{5,\iota}\left|r_{t-\iota}\right|+\beta_6\frac{\widehat{\sigma_{d(t)}}}{\sqrt{24}}+\sum_{h=19}^{23}\beta_{7,h}Dum_{FRI_h,t}+\varepsilon_t.$$

absolute announcement surprises provide a statistic of 0.0199.[12] Thus, the announcement surprises are fairly important but not as important as some other features of the data, confirming the views of Andersen and Bollerslev (1998).

Figure 1 illustrates the predictive power of various components of regression (2) by showing the average actual volatility over various hours of the week along with average predicted volatility for those hours. The periodic component shows the greatest covariation with actual volatility but the announcement predictors and the lagged returns also help explain the average actual volatility.

Table 2 shows the estimated regression coefficients and the t-statistics from equation (2). Most—but not all—of the news surprise coefficients are positive, indicating that larger surprises increase volatility. Some of the news surprise coefficients are perverse (negative), which often results from their correlation with the periodic components and/or the announcement indicators. Of all the German/euro announcements, only German real GDP growth is significant and positive. The U.S. announcement indicator is significant, whereas the German/euro indicator is essentially zero—that is, U.S. announcements raise volatility but German announcements do not. The significance of the U.S. announcement indicator confirms the results of Andersen et al. (2003), who use high-frequency (5-minute) data from 1992 through 1998 to study the effects of a large set of U.S. and German announcements on the conditional mean and the conditional volatility of DEM/USD, USD/GBP (British pound sterling), JPY/USD, CHF (Swiss franc)/USD, and USD/EUR exchange rates. The authors find that

[12] The addition of indicator variables for the Friday evening hours also improves the fit of the model. The intraday periodic variables do not fit these weekly fluctuations.

Table 2

Regression Coefficients from Equation (2)

Independent variable	Coefficient	*t*-Statistic
U.S. announcement dummy	0.025	10.012*
German/Euro announcement dummy	0.000	–0.084
U.S.: Real GDP: Advance	0.050	4.160*
U.S.: Real GDP: Preliminary	–0.014	–1.179
U.S.: Real GDP: Final	0.011	0.963
U.S.: Business inventories	–0.002	–0.287
U.S.: Capacity utilization rate: Total industry	–0.036	–3.051†
U.S.: Consumer confidence	0.036	5.177*
U.S.: Construction spending	0.047	5.676*
U.S.: CPI	0.004	0.553
U.S.: Consumer credit	–0.027	–3.841†
U.S.: New orders: Advance durable goods	0.006	0.829
U.S.: New orders	–0.017	–2.438†
U.S.: Housing starts	–0.014	–1.936
U.S.: Industrial production	0.032	2.736*
U.S.: Composite Index of Leading Indicators	–0.012	–1.774
U.S.: ISM: Manufacturing Composite Index	0.045	5.118*
U.S.: Employees on nonfarm payrolls	0.173	25.861*
U.S.: New home sales	0.004	0.609
U.S.: PCE	–0.016	–2.123†
U.S.: Personal income	–0.014	–1.930
U.S.: PPI	–0.009	–1.200
U.S.: Retail sales	0.021	2.208*
U.S.: Retail sales ex motor vehicles	0.015	1.497
U.S.: Trade balance: Goods & services (BOP)	0.050	7.131*
U.S.: Government surplus/deficit	–0.010	–1.365
U.S.: Initial unemployment claims	0.001	0.279
Euro area: CPI flash estimate Yr/Yr %Chg	0.005	0.768
Euro area: IP WDA Yr/Yr %Chg	0.013	1.868
Euro area: Money supply M3 Yr/Yr %Chg	–0.012	–1.581
Euro area: Harmonized CPI Yr/Yr %Chg	0.002	0.349
Euro area: Unemployment rate	0.006	0.865
Euro area: PPI Yr/Yr %Chg	0.006	0.817
Euro area: Retail sales WDA Yr/Yr %Chg	–0.010	–1.422
Euro area: Trade balance	–0.005	–0.646
Euro area: Preliminary real GDP Yr/Yr %Chg	-0.018	–1.527
Euro area: Final real GDP Yr/Yr %Chg	–0.003	–0.276

Table 2, cont'd

Regression Coefficients from Equation (2)

Independent variable	Coefficient	*t*-Statistic
Germany: Current account balance	–0.011	–1.046
Germany: Final cost of living	0.001	0.094
Germany: Preliminary cost of living	–0.018	–2.486†
Germany: IP: Total industry Mo/Mo %Chg	–0.007	–1.061
Germany: Current account balance	–0.005	–0.778
Germany: PPI: Mfg. Yr/Yr %Chg	0.002	0.327
Germany: Real retail sales Yr/Yr %Chg	0.001	0.163
Germany: Current account: Trade balance	0.020	1.909
Germany: Real GDP Qtr/Qtr %Chg	0.042	3.493*
Cos_q1	–0.011	–22.020†
Cos_q2	0.004	8.137*
Cos_q3	–0.009	–20.889†
Cos_q4	–0.007	–15.360†
Sin_q1	0.016	35.591*
Sin_q2	–0.001	–3.291†
Sin_q3	0.003	6.110*
Sin_q4	–0.007	–14.930†
Absolute return lag1	0.095	21.428*
Absolute return lag2	0.043	9.632*
Absolute return lag3	0.024	5.418*
Absolute return lag4	0.031	6.923*
Absolute return lag5	0.022	5.019*
Constant	–0.009	–8.002†
GARCH daily volatility	3.010	47.135*
Friday_7 p.m.	–0.039	–11.173†
Friday_8 p.m.	–0.045	–12.959†
Friday_9 p.m.	–0.043	–12.431†
Friday_10 p.m.	–0.033	–9.408†
Friday_11 p.m.	–0.026	–7.412†

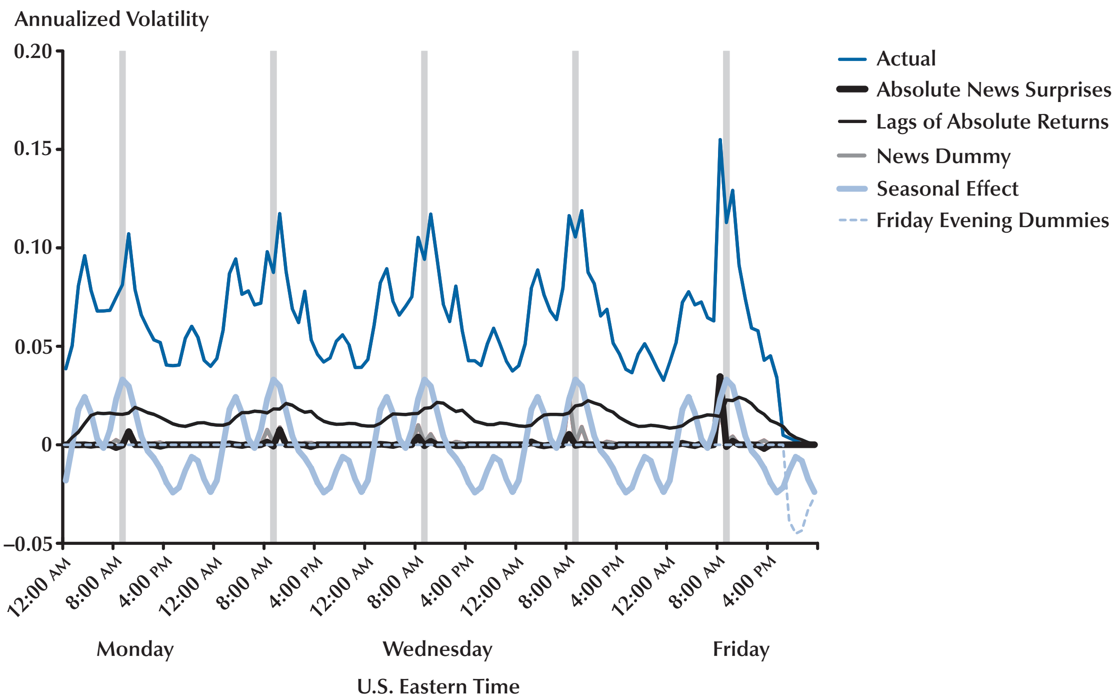
NOTE: The table shows the regression coefficients from estimating equation (2) (below) on absolute USD/EUR log changes, over the sample period November 5, 2001, to March 12, 2010. BOP, balance of payments; CPI, consumer price index; GDP, gross domestic product; IP, industrial production; ISM, Institute for Supply Management; PCE, personal consumption expenditures; PPI, producer price index; WDA, work days adjusted. *Statistically significant positive coefficients; †, statistically significant negative coefficients.

$$|r_t| = \alpha + \beta_{1,US} Dum_{USann,t} + \beta_{1,for} Dum_{forann,t} + \sum_{j=1}^{N} \beta_{2,j} |s_{j,t}| + \sum_{q=1}^{4} \left(\beta_{3,q} \cos\left(\frac{q2\pi t}{24}\right) + \beta_{4,q} \sin\left(\frac{q2\pi t}{24}\right) \right)$$

$$+ \sum_{\iota=1}^{5} \beta_{5,\iota} |r_{t-\iota}| + \beta_6 \frac{\widehat{\sigma_{d(t)}}}{\sqrt{24}} + \sum_{h=19}^{23} \beta_{7,h} Dum_{FRI_h,t} + \varepsilon_t .$$

Figure 1

Average Actual and Predicted Hourly Volatility for the USD/EUR



NOTE: The figure shows the average actual and predicted volatility of USD/EUR absolute log changes, estimated with equation (2) over the sample period November 5, 2001, to March 12, 2010. See equation (2) below.

$$|r_t| = \alpha + \beta_{1,US} Dum_{USann,t} + \beta_{1,for} Dum_{forann,t} + \sum_{j=1}^{N} \beta_{2,j} |s_{j,t}| + \sum_{q=1}^{4} \left(\beta_{3,q} \cos\left(\frac{q2\pi t}{24}\right) + \beta_{4,q} \sin\left(\frac{q2\pi t}{24}\right) \right)$$
$$+ \sum_{\iota=1}^{5} \beta_{5,\iota} |r_{t-\iota}| + \beta_6 \frac{\widehat{\sigma_{d(t)}}}{\sqrt{24}} + \sum_{h=19}^{23} \beta_{7,h} Dum_{FRI_h,t} + \varepsilon_t .$$

both the magnitude of the surprise and the pure announcement effect are significant.[13]

In summary, the results in Table 1 indicate that Andersen and Bollerslev (1998) were correct to argue that announcements are important explanatory variables for volatility, though not as important as intraday periodicity and daily volatility. Likewise, Table 2 confirms the findings of Ederington and Lee (1993) that U.S. nonfarm payroll and U.S. trade balance surprises are among the most important for volatility.

Volatility and News Arrival

Not all news consists of macro announcements. Information about the international economy and politics arrives continuously in financial markets via newswire reports. The literature on the impact of information on stock trading and volatility (i.e., Berry and Howe, 1994, and Mitchell and Mulherin, 1994) helped motivate research in the foreign exchange market on whether such

[13] Andersen et al. (2007) use a similar model to study the effects of macroeconomic news releases on asset returns across countries and over the business cycle. They find evidence that news creates asset price jumps in all markets. They also use macro release indicators to model conditional volatility but do not focus on those results.

public information flow affects market volume and volatility.[14]

Most papers documenting the impact of information arrival use some measure of the frequency of headlines from wire service news agencies such as Reuters. DeGennaro and Shrieves (1997), however, incorporate unexpected quote arrival as a proxy for information arrival.[15] This strategy is implicitly endorsed by Melvin and Yin (2000), who show that public information arrival influences both quote frequency and GARCH volatility of high-frequency JPY/USD and DEM/USD data.

The most common theme in this literature is that information arrival typically *does* increase volatility (DeGennaro and Shrieves, 1997; Eddelbüttel and McCurdy, 1998; Joines, Kendall, and Kretzmer, 1998; Melvin and Yin, 2000; Chang and Taylor, 2003). Melvin and Yin (2000) interpret this result as casting doubt on proposals to apply "sand-in-the-wheels" transaction taxes that would reduce allegedly self-generated foreign exchange volatility.

There are exceptions to the rule that news arrival boosts volatility, however. DeGennaro and Shrieves (1997) find that unscheduled announcements actually reduce volatility for 20 minutes, perhaps inducing traders to pause to consider unexpected information. And not all news is created equal. Chang and Taylor (2003) find that Bundesbank news is most significant for DEM/USD volatility, and major U.S. and German announcements are more significant than simple headline counts.

Eddelbüttel and McCurdy (1998) use Reuters' news headlines as a proxy for news arrival and confirm that the addition of such a news variable renders the GARCH-implied variance process much less persistent. This fact appears to confirm the intuitively attractive proposition that persistence in news arrival drives part of the volatility persistence captured by GARCH models.

The literature also shows, however, that public information arrival cannot explain the entire increase in volatility. Joines, Kendall, and Kretzmer (1998) and Chang and Taylor (2003) argue that trading must also release private information that hikes volatility. Researchers working with order flow data would further explore this point.

Volatility and Non-U.S. Announcements

The earliest papers on announcement effects studied the effects of U.S. announcements almost exclusively, but researchers soon began to consider how announcements from a variety of countries influence foreign exchange volatility. Many of these studies used variations on the popular GARCH model, including the EGARCH-in-mean (exponential GARCH-in-mean) model (Kim, 1998, 1999), trivariate GARCH to compare announcement effects on foreign exchange rates and Italian bond markets (Fornari et al., 2002), and FIGARCH (fractionally integrated GARCH) to account for possible long memory (Han, 2004). Other studies look at the effect of the announcement itself versus the information content (Kim, McKenzie, and Faff, 2004), the effect of conflicting information (Laakkonen, 2004) or heterogeneous information (Hashimoto and Ito, 2009), and asymmetric responses to news (Han, 2004).

These papers frequently contain two themes. First, most studies find that U.S. news has a greater impact on volatility than foreign news (e.g., Cai, Joo, and Zhang, 2009; Evans and Speight, 2010; Harada and Watanabe, 2009); however, Kopecký (2004) is an exception in finding that Czech announcements raise CZK (Czech crown [koruna])/USD volatility but—very curiously—U.S. announcements do not. The second common theme is that the volatility effect of announcements potentially depends on many factors: heterogeneous expectations, conflicting information, the source of the shocks, the sign of the shock, and whether the announcement is scheduled or unscheduled.

[14] Public information flow is effectively synonymous with *news arrival*, which refers to the rate at which news headlines or quotes are observed rather than the outcome of specific announcements.
Chaboud, Chernenko, and Wright (2008) introduce a new dataset of volume in foreign exchange markets from the Swiss Electronic Bourse system. Although they do not study volatility specifically, they find that volume increases after U.S. macroeconomic announcements regardless of whether the announcement is expected or unexpected. For unexpected news, a price jump precedes the increase in volume.

[15] Financial traders receive electronic feeds that allow them to see quotes on asset prices. *Quote arrival* is the rate at which such quotes are updated. *Unexpected quote arrival* is the surprise component of this measure.

Announcements and Jumps

Researchers have noted *jumps*—discontinuities in asset prices—for some time. The efficient markets hypothesis easily explains many jumps because it predicts very rapid systematic price reactions to news surprises to prevent risk-adjusted profit opportunities. Decomposing volatility into jumps and time-varying diffusion volatility is important because these two components have different implications for modeling, forecasting, and hedging. For example, persistent time-varying diffusion volatility would help forecast future volatility, while jumps might contain no predictive information or even distort volatility forecasts (Neely, 1999, and Andersen, Bollerslev, and Diebold, 2007). Therefore, it makes sense to investigate the effect of announcements on jumps.

Goodhart et al. (1993) first suggested the importance of accounting for news-induced discontinuities in modeling exchange rates. The authors study the effect of announcements on the time-series properties of exchange rates using a 3-month sample (April 9 to July 3, 1989) of high-frequency USD/GBP data from Reuters. The authors make strong claims that including news indicators in the conditional mean and variance equations of a GARCH-in-mean (GARCH-M) model renders both of these processes stationary. This is similar to the well-known phenomenon that discontinuities in macro series lead to spurious findings of non-stationarity (Perron, 1990).[16] At high frequencies, conditional volatility appears to be very persistent; accounting for shocks to conditional volatility greatly reduces this persistence.

To link jumps to economic news, Johnson and Schneeweis (1994) introduce an announcement effect parameter to Jorion's (1988) jump-diffusion model, permitting the conditional variance to depend on an announcement indicator. Using daily data between 1988 and 1990, the authors relate jumps in the JPY, GBP, and DEM exchange rates to four announcements from U.S., British, German, and Japanese sources. They find that certain real announcements—U.S. trade balance and industrial production news—cause larger volatility movements than do money supply and inflation news. U.S. news influences currency market variance more than does foreign news, and covariances between the exchange rates were highest on U.S. announcement days. Conditional variance and jump-diffusion models outperform simple diffusion and homoskedastic models. Incorporating news indicators in a diffusion model fits the conditional variance process better than estimating a jump process.

Fair (2003) turns the usual procedure for examining the relation between announcements and large exchange rate changes on its head.[17] Instead of estimating a jump model of exchange rates that incorporates macro surprises, Fair looks for the largest changes in U.S. foreign exchange (and stock and bond) futures tick prices from 1982 to 2000 and then relates those changes to contemporaneous news. Monetary, price level, employment, and trade balance news are often associated with large changes in U.S. foreign exchange futures prices.

Advances in econometric jump modeling enabled later researchers to better examine the relation between announcements and jumps. Specifically, Barndorff-Nielsen and Shephard's (2004) bipower procedure enabled researchers to pinpoint the dates and magnitudes of exchange rate jumps without needing to specify a likelihood function. Barndorff-Nielsen and Shephard (2004) observe that many jumps appear to correspond to macroeconomic releases, which is consistent with Andersen et al. (2003, 2007).

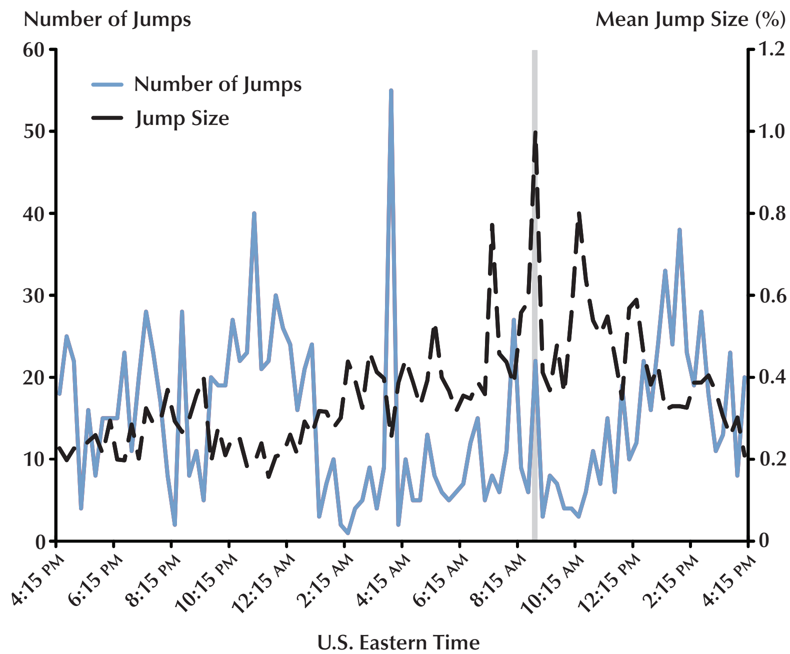
The Barndorff-Nielsen and Shephard (2004) bipower procedure estimates the sum of jumps during a period, usually a day. It does not pin down the precise times of those jumps, however, which makes it difficult to precisely link jumps to events such as news releases. Lee and Mykland (2008) developed another jump-detection method that compares each return, standardized by local volatility, with the distribution of the maximal diffusion return over the sample. The Lee and

[16] While it is plausible that accounting for discontinuities would render the conditional variance much less persistent, a broader view of the data indicates that nominal exchange rates are very unlikely to be stationary—and one cannot draw conclusions about such behavior from three months of data in any case.

[17] Andersen and Bollerslev (1998) perform a similar exercise, examining whether any obvious political or economic events could explain the 25 largest 5-minute returns in their sample.

Figure 2

Number of Significant Jumps and Mean of Absolute Jumps Conditional on the Intraday Period



NOTE: The *x*-axis represents intraday time (U.S. ET). The left *y*-axis displays the number of significant jumps ($\alpha = 0.1$), while the right *y*-axis shows the mean absolute value of significant jumps in the USD/EUR exchange rate. The solid line denotes the number of jumps and the dashed line denotes mean jump size. The vertical gray line denotes the interval containing 8:30 a.m., the time of most news arrivals. The sample period is 1987-2004.

SOURCE: From Figure 2 in Lehaye, Laurent, and Neely (2010).

Mykland (2008) method permits one to more precisely time jumps than does bipower variation.

Lahaye, Laurent, and Neely (2010) use the Lee and Mykland (2008) technique to determine that U.S. macro announcements explain jumps and cojumps—simultaneous jumps in multiple markets—across equity, bond, and foreign exchange markets.[18] Nonfarm payroll and federal funds target announcements are the most important news across asset classes, while trade balance shocks are also important for foreign exchange jumps.

[18] Beine et al. (2007) use macro announcements as control variables in a study of the effects of U.S., German, and Japanese foreign exchange intervention on the continuous and discontinuous components of DEM-EUR/USD and JPY/USD exchange rate volatility. They estimate exchange rate jumps with bipower variation.

Figure 2 illustrates the frequency and size of shocks in the USD/EUR market by time of day. Exchange rate jumps are more frequent around 8:30 a.m., 4 p.m. to 8 p.m., and 10 p.m. to 2 a.m. U.S. ET. The largest jumps occur at the times of major macro news; smaller liquidity jumps are associated with periods of low volatility (i.e., Tokyo lunch and early Asian trading).

Lahaye, Laurent, and Neely (2010) use tobit-GARCH and probit models to formally examine the relation between U.S. news and a variety of asset price jumps and cojumps, respectively. Table 3 shows that the tobit-GARCH regression formally confirms that nonfarm payroll (NFP), federal funds target announcements, trade balance reports, preliminary GDP, government fiscal announcements, and consumer confidence

Table 3

Tobit-GARCH Models for Jumps

Variable	USD/EUR coefficient	$p > \|t\|$	JPY/USD coefficient	$p > \|t\|$	USD/GBP coefficient	$p > \|t\|$	CHF/USD coefficient	$p > \|t\|$
Consumer confidence			0.74	0.00	0.38	0.08	0.43	0.02
Consumer credit					0.06	0.99	–0.13	0.99
CPI	0.01	1.00	0.09	0.99			–0.06	0.99
Federal funds target	0.88	0.00	0.72	0.00	0.66	0.00	0.57	0.00
Advanced GDP					0.40	0.83	0.48	0.81
Preliminary GDP	0.81	0.00	0.84	0.01			0.58	0.04
Government fiscal surplus/deficit	–0.55	0.17	–0.72	0.08	–0.32	0.66	–0.62	0.08
Manufacturing index	0.24	0.81	–0.21	1.00	–0.04	1.00	0.54	0.12
Nonfarm payroll	0.98	0.00	0.35	0.25	0.16	0.94	0.43	0.00
PPI	–0.70	0.99	–0.82	0.99	–0.15	0.58	–1.02	0.67
Retail sales	–0.21	0.99					–1.18	0.99
Trade balance	0.43	0.05	0.02	1.00	0.17	0.89	0.47	0.02
Omega	0.30	0.00	0.26	0.00	0.25	0.00	0.52	0.00
Alpha1	0.19	0.00	0.18	0.00	0.28	0.00	0.26	0.00
Alpha2							0.09	0.00
Beta	0.49	0.00	0.60	0.00	0.28	0.00		
Function value	–7090.68		–7542.77		–7727.96		–7331.87	
No. of observations	352,127		351,359		352,799		352,319	

NOTE: The latent tobit jump variable is denoted by $Jump^*_{t,i} = \mu + \eta_{t,i} + \mu_{t,i} + \xi_{t,i} + \varepsilon_{t,i}$, where $|Jump_{t,i}| = Jump^*_{t,i}$ if $Jump^*_{t,i} > 0$ and $|Jump_{t,i}| = 0$ if $Jump^*_{t,i} \leq 0$; $\varepsilon_{t,i}|I_{t,i-1}$ is $N(0,\sigma_t^2)$. The variance σ_t^2 is assumed to follow an ARCH or GARCH process. $|Jump_{t,i}|$ represents significant jumps at the 10 percent level. $\eta_{t,i}$ controls for day of the week effects (not reported) and $\mu_{t,i}$ includes absolute surprises concerning macro announcements. $\xi_{t,i}$ controls for intradaily periodicity (not reported). Estimates and robust *p*-values ($2x(1 - \text{Prob}(X < |tstat|))$) are reported for surprise coefficients that are significant at the 10 percent level in at least one series, as well as the ARCH and GARCH coefficients, where *X* is a *t*-distributed random variable with N–K (no. of observations – no. of parameters) degrees of freedom under the null and *tstat* is the estimated coefficient over its standard error. Regressors with no contemporaneous match with significant jumps are excluded from the model. Function value is the maximized log-likelihood function value. The exchange rate samples start in January 1990 and end on October 1, 2004. CHF, Swiss franc; CPI, consumer price index; EUR, euro; GBP, British pound sterling; GDP, gross domestic product; JPY, Japanese yen; PPI, producer price index; USD, U.S. dollar.

SOURCE: From Table 6 in Lahaye, Laurent, and Neely (2010).

Table 4

Probit Models for Cojumps

Variable	USD/EUR-USD/GBP coefficient	$p > \|t\|$	USD/EUR-JPY/USD coefficient	$p > \|t\|$	USD/EUR-CHF/USD coefficient	$p > \|t\|$	USD/GBP-JPY/USD coefficient	$p > \|t\|$	USD/GBP-CHF/USD coefficient	$p > \|t\|$	JPY/USD-CHF/USD coefficient	$p > \|t\|$
Construction spending					–7.41	0						
Consumer confidence											0.73	0
Federal funds target	1.08	0	0.86	0	0.83	0	0.90	0	0.89	0	0.74	0.01
Preliminary GDP			0.87	0	0.6	0.02					0.83	0
Government fiscal surplus/deficit			0.23	0.07	0.17	0.18					0.25	0.05
Manufacturing index					1.50	0						
Nonfarm payroll			0.65	0	0.79	0					0.61	0
Trade balance					0.76	0.02						
Function value	–1842.90		–1181.87		–3130.59		–742.60		–1610.76		–933.24	
Pseudo-R^2	0.04		0.04		0.03		0.05		0.04		0.05	
No. of observations	349,355		348,967		349,557		348,593		349,542		348,619	

NOTE: The latent probit cojump variable is denoted by $CO\ Jump^*_{t,i} = \mu + \eta_{t,i} + \mu_{t,i} + \xi_{t,i} + \varepsilon_{t,i}$, where $CO\ Jump_{t,i} = 1$ if $CO\ Jump^*_{t,i} > 0$ and $CO\ Jump_{t,i} = 0$ if $CO\ Jump^*_{t,i} \leq 0$. $\varepsilon_{t,i}$ is $NID(0,1)$. $CO\ Jump_{t,i}$ is the cojump (simultaneous significant jumps) indicator. $\eta_{t,i}$ controls for day of the week effects (not reported) and $\mu_{t,i}$ includes absolute surprises concerning macro announcements. $\xi_{t,i}$ controls for intradaily seasonality (not reported). Estimates and robust p-values ($2x(1 - \text{Prob}(X < |tstat|))$) are reported for surprise coefficients that are significant at the 10 percent level in at least one series, as well as the ARCH and GARCH coefficients, where X is a t-distributed random variable with N–K (no. of observations – no. of parameters) degrees of freedom under the null and *tstat* is the estimated coefficient over its standard error. Regressors with no contemporaneous match with significant cojumps are excluded from the model. We further report the maximized log likelihood function value, and the McFadden R^2, which is $1 - (\text{LogLik}_1/\text{LogLik}_0)$ (i.e., 1 minus the ratio of the log-likelihood function value of the full model to the constant-only model). The exchange rate samples start in January 1990 and end on October 1, 2004. CHF, Swiss franc; EUR, euro; GBP, British pound sterling; GDP, gross domestic product; JPY, Japanese yen; USD, U.S. dollar.

SOURCE: From Table 7 in Lahaye, Laurent, and Neely (2010).

surprises contribute to foreign exchange jumps. Table 4 likewise shows that a probit model consistently and strongly links cojumps to macro surprises, such as those to the federal funds rate target, NFP, and preliminary GDP. It is noteworthy that federal funds target surprises significantly explain cojumps in every currency pair. In summary, research has shown that many announcements cause jumps and cojumps and that a substantial proportion of jumps are associated with announcements.

Order Flows and Foreign Exchange Volatility

News might create order flows—signed transaction flows—that transmit private information to the foreign exchange market. Private agents combine public news releases with their own private information, and their publicly observable decisions may convey that private information.[19] For example, a business might observe an uptick in industrial production, revise its estimates of future demand accordingly, and decide to build a new plant—but only if the firm's privately known cost structures would make it expect to profit from that decision. If news announcements cause the release of private information that generates conflicting trades, then this provides a channel through which news can affect volatility over a prolonged period.

Because obtaining order flow data is expensive and/or difficult, some researchers have used proxies for order flow: Cai et al. (2001) use yen positions held by major market participants, and Bauwens, Ben Omrane, and Giot (2005) use quote frequency. Most researchers have used data from electronic brokers such as Reuters D2000-1 (Evans, 2002), Reuters D2000-2 (Dominguez and Panthaki, 2006, and Carlson and Lo, 2006), or Electronic Brokerage Systems (Berger, Chaboud, and Hjalmarsson, 2009). Others have used proprietary datasets from commercial banks (Savaser, 2006, and Frömmel, Mende, and Menkhoff, 2008). Unfortunately, the difficulty of obtaining long spans of order flow data has left many of the studies of announcements and order flow with samples only a few months long. This limitation has prevented those studies from drawing clear conclusions about the effect of specific announcements on order flow.

The main finding from the literature on order flow and announcements is that news releases public information that immediately impacts prices and volatility and impacts volume through order flow with a delay. The release of public information causes an immediate "average" effect on prices, as well as delayed trading based on both the news and private information (Evans and Lyons, 2005). This delayed trading produces the protracted volatility found in the literature. In fact, the indirect impact of news on volatility through order flow is more important than the direct impact of the news itself (Cai et al., 2001). Likewise, Evans (2002) estimates some fairly complex microstructure models that decompose macro news (and other shocks) into common knowledge and non-common knowledge shocks. Evans (2002) argues that non-common knowledge shocks are of greater importance than textbook models emphasize.

The delayed effects of order flow can contribute to volatility for hours after announcements, particularly if the announcement is important and unscheduled. Carlson and Lo (2006) examine the reaction of the Reuters D2000-2 electronic order book on foreign exchange transactions to a single announcement—an October 9, 1997, surprise interest rate hike by the Bundesbank, aimed at heading off inflation pressures. Volatility remained high for about 2 hours after this unscheduled and surprising news. There were also price jumps after the announcement: 14 of the 19 largest price changes in a 4-day window occurred within 2 hours after the release.

It is possible, of course, that volatility persists after news is released either because of persistence

[19] Hasbrouck (1991) reasons that news surprises should not directly affect order flow under rational expectations because although news might cause an immediate price jump to a new equilibrium, it should not cause systematic orders—or the price effects from those predictable orders would themselves be predictable, creating a profit opportunity. Although the Hasbrouck reasoning has strongly influenced the microstructure literature, Evans (2010) lays out two microstructure models in which such reasoning fails because announcements can affect order flow through dealers' risk management practices. Dealers alter their quotes to produce predictable patterns in order flow to better manage their inventory risk.

in news/order flow or persistence in sensitivity to news/order flow. Berger, Chaboud, and Hjalmarsson (2009) tackle the difficult problem of disentangling the importance of these two effects. Using six years of high-frequency exchange rate data, Electronic Broking Services (EBS) order flows, and news, they conclude that both factors contribute to the persistence of volatility.

The theoretical and empirical microstructure literature has found that much of the effect of order flow consists of transmitting private information to markets. The amount of information depends on the type of order flow. Financial customers are thought to have better information on asset prices from their own trading and research, whereas commercial firms are considered to be price takers that trade to import or export goods rather than because the firms' agents think that they have superior information about future asset prices. That is, the *type* of order flow matters. Frömmel, Mende, and Menkhoff (2008) find that only order flow from banks and financial customers (i.e., informed order flow) is linked to higher foreign exchange volatility.[20] Savaser (2006) finds that investors—probably informed traders—substantially increase their use of limit orders—stop-loss and take-profit orders—prior to news releases and that accounting for this surge substantially improves the ability to explain the exchange rate jumps that follow news.

Perhaps more surprising than the post-announcement increase in volatility is the fact that informed trading can apparently increase volatility *before* announcements as the informed traders take speculative positions based on their private information (Bauwens, Ben Omrane, and Giot, 2005).

Not only does the type of order flow matter, but the definition of "news" matters as well. Dominguez and Panthaki (2006) argue for expanding the definition of news to include both "fundamental" and "non-fundamental" news. Non-fundamental news includes technical analysis indicators, political news, and important private sector changes, such as mergers and acquisitions. The authors suggest taking a broader view of relevant variables in models of exchange rate determination.

In summary, the literature has found that (i) orders and order flow often respond to news, (ii) informed order flow has greater effects, and (iii) persistence in order flow and persistence in sensitivity to order flow both produce persistence in volatility in the wake of many announcements.

Recent Research on Monetary Policy Announcements and Exchange Rate Volatility

Several developments in central banking led researchers to renew attention to the effects of monetary policy announcements in the late 1990s. First, the Bank of England gained operational independence in the conduct of monetary policy from the government of the United Kingdom in 1998.[21] Second, the European Central Bank (ECB) began to conduct a common monetary policy for the European Monetary Union as of January 1, 1999.[22] Third, policymakers and researchers began to seriously reconsider the importance of communication in the 1990s and central banks responded by publicly explaining their policy actions.[23] These policies prompted economists to begin to reconsider the effects of monetary structure, policy actions, and communications on asset prices and volatility.

The two most common themes of research on the effects of monetary policy news are as fol-

[20] *Informed* order flow would be order flow that is generated by private information and speculates on a change in asset prices. In contrast, *uninformed* order flow would be generated by demands for commercial or hedging purposes and would not be predicated on private information that informs expectations of changes in asset prices.

[21] On May 6, 1997, Chancellor of the Exchequer Gordon Brown announced that the government of the United Kingdom would grant the Bank of England operational independence over monetary policy. The Bank of England Act 1998 formalized this arrangement.

[22] The original members of the European Monetary Union were Austria, Belgium, Finland, France, Germany, Ireland, Italy, Luxembourg, the Netherlands, Portugal, and Spain.

[23] For example, the Federal Open Market Committee (FOMC) began to contemporaneously announce policy actions in 1994 and adopted this as formal policy in 1995. Starting in August 1997, each FOMC policy directive has included the quantitative value of the "intended federal funds rate." And since 1999, the FOMC has issued a press release after each meeting with the value for the "intended federal funds rate" and, in most cases, an assessment of the balance of risks (Poole, Rasche, and Thornton, 2002).

lows: (i) that surprising policy actions, such as changes in interest rates or currency parities, increase volatility and (ii) that clarification of longer-term policy reduces volatility. This seems to be true of the ECB, the Bank of England, and the Bank of Canada. Using a Markov-switching model, Sager and Taylor (2004) find that volatility tends to increase after an ECB interest rate announcement, peaking 15 minutes later but remaining elevated for an hour. Conrad and Lamla (2010) likewise show that the ECB's interest rate decision and press conference strongly affect EUR/USD volatility but its later question and answer session produces no substantial effect. Using data from 1997 to 2007, Melvin et al. (2009) find that the USD/GBP Markov volatility-generating process changes entirely after surprising Bank of England interest rate announcements. Hayo and Neuenkirch (2009) use daily GARCH-M models to determine that Canadian interest rate changes raise CAD (Canadian dollar)/EUR volatility. The close relationship between monetary policy and exchange rate policy means that a change in exchange rate parities implies a change in monetary policy analogous to a change in the expected interest rate path. So it should not be surprising that Chelley-Steeley and Tsorakidis (2009) find that the devaluation of the Greek drachma increased exchange rate volatility. Interest rate changes and changes in currency parity are not the only central bank actions that can change foreign exchange volatility. Jansen and De Haan (2005) find that indicators of statements from ECB and national central bank officials raise USD/EUR EGARCH volatility.

Central bank actions that fix expectations about future policy without changing current policy often reduce volatility, however. Bank of Canada communications lower CAD/EUR volatility (Hayo and Neuenkirch, 2009), and Greece's announcements that it would be joining the European Exchange Rate Mechanism and its commitment to the euro zone reduced exchange rate volatility (Chelley-Steeley and Tsorakidis, 2009).

Recent Research on State-Dependent Reactions

Economists have considered how the response of volatility to news announcements might depend on the nature of the economy or the nature of the news. Pearce and Solakoglu (2007) reject asymmetry and nonlinearity in DEM/USD and JPY/USD volatility reactions but find some evidence of changes across the state of the business cycle.

Motivated by the idea that investors should react more strongly to high-quality information, Laakkonen and Lanne (2009) use 6 years of high-frequency USD/EUR data and 20 announcements to find that "precise" U.S. news announcements affect volatility more than imprecise announcements. The authors measure precision as the degree to which the previous month's news announcements are not revised.

DISCUSSION AND CONCLUSION

This article has reviewed the literature on how news affects foreign exchange volatility. The ability to understand and quantify asset price uncertainty is crucial to managing risk and choosing portfolio composition.

The research on announcements and volatility has been particularly useful because it highlights the role of announcements in contributing to two of the main characteristics of volatility: periodicity and jumps. The most basic result of the literature is that trading and volatility typically increase for about an hour after an announcement. The same announcements that strongly affect foreign exchange returns—nonfarm payrolls, trade balance, advance GDP, and interest rate changes—also tend to increase volatility (Ederington and Lee, 1993).

Disentangling the contributions of macroeconomic news from those of other periodic market effects—such as market openings and closings—challenged the early authors in this literature (Payne, 1996, and Andersen and Bollerslev, 1998). Indeed, studies of intraday and intraweek periodicity in volatility motivated researchers to consider announcement effects on volatility.

Scheduled macroeconomic announcements shed light on market microstructure because they provide a natural experiment through which to study the release of public information on volatility. A series of studies established that public information flow affects market volume and volatility (Ederington and Lee, 2001; Melvin and Yin, 2000; Chang and Taylor, 2003). Other studies have used microstructure theory to motivate investigations into how volatility might depend on the precision of the information in the news, the state of the business cycle, and/or the heterogeneity of beliefs (Baillie and Bollerslev, 1991, and Kim, 1998).

Although the first studies of news volatility effects used U.S. news reports and USD exchange rates, later studies branched out to study the effect of foreign news and broader definitions of news. Most such work has found that U.S. news has stronger effects on foreign exchange volatility than does foreign news (Cai, Joo, and Zhang, 2009; Evans and Speight, 2010; Harada and Watanabe, 2009).

Announcements frequently cause jumps, which are an important part of foreign exchange volatility (Goodhart et al., 1993; Fair, 2003; Andersen et al., 2003; Lahaye, Laurent, and Neely, 2010). The development of better tests for price discontinuities has aided the more recent jump studies. Removing such jumps from the volatility process improves autoregressive volatility forecasts (Neely, 1999, and Andersen, Bollerslev, and Diebold, 2007).

More recently, researchers have established that news has a prolonged effect on order flow, releasing private information, which leads to sustained increases in volatility (Cai et al., 2001; Evans, 2002; Evans and Lyons, 2005; Frömmel, Mende, and Menkhoff, 2008). Berger, Chaboud, and Hjalmarsson (2009) show that time variation in sensitivity to order flow contributes to the persistence of volatility. And the definition of "news" has expanded over time (see Dominguez and Panthaki, 2006).

Monetary policy communications can raise foreign exchange volatility when they describe a surprising change in current interest rates, but they tend to lower volatility when the communications anchor longer-term expectations of policy (Sager and Taylor, 2004; Melvin et al., 2009; Conrad and Lamla, 2010).

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APPENDIX

Summary of the Literature on Estimating Announcement Effects on the Conditional Volatility and Jumps of Exchange Rate Returns

Reference	Abstract/Description*
Madura and Tucker (1992)	We investigate the effects of U.S. balance of trade deficit announcements on the ex ante volatility of foreign exchange rates. Specifically, we analyze the association between currency option implied standard deviation (ISDs) and the surprise component of monthly merchandise trade deficit disclosures. Our results indicate that larger surprises, regardless of their sign, are associated with increased currency ISDs. We also find that deficit disclosures regardless of their content, temper market uncertainty on average. Finally we find that larger than expected deficits tend to depreciate the U.S. dollar.
Goodhart et al. (1993)	This paper uses an extremely high frequency data set on the dollar-sterling exchange rate to investigate the impact of news events on the very short-term movements in exchange rates. The data set is a continuous record of the quoted price for the exchange rate on the Reuters screen. As such it records some 130,000 observations over an 8-week period. The paper investigates the time-series properties of the data using orthodox regression models, and then by making allowance for a time-varying conditional variance. The conclusions vary significantly in moving to this more sophisticated model. The exercises are repeated now incorporating news announcement effects, letting these affect the level of the exchange rate and then the conditional variance process. Again it is found that the conclusions are radically altered in moving to the increasingly sophisticated model.
Ederington and Lee (1993)	[A]nnouncements are responsible for most of the observed time-of-day and day-of-the-week volatility patterns in these markets. While the bulk of the price adjustment to a major announcement occurs within the first minute, volatility remains substantially higher than normal for roughly fifteen minutes and slightly elevated for several hours.
Hogan and Melvin (1994)	We examine the role that news and heterogeneous expectations play in the persistence of exchange rate volatility, or so-called “meteor shower” effects. Our empirical focus is on the U.S. trade balance news, which is shown to have a significant and persisting effect on the exchange rate and its conditional variance. Furthermore, the impact of U.S. trade balance “news” is not isolated to the U.S foreign exchange market. The degree to which U.S. trade balance “news” affects other geographical market locations is functionally related to heterogeneous priors.

*Excerpts are directly quoted from the original sources.

NOTE: The following general abbreviations are used in the appendix table: AAC, average absolute value of price changes; ARCH, autoregressive conditional heteroskedasticity; AR-FIGARCH; autoregressive FIGARCH; EGARCH, exponential GARCH; FAC, first-order autocorrelation of log price changes; FIGARCH, fractionally integrated GARCH; FX, foreign exchange; GARCH, generalized autoregressive conditional heteroskedasticity; GARCH-M, GARCH-in-mean; NP, number of prices; PR, price fluctuation range; SD, standard deviation. Unless stated otherwise, announcements are for the United States. The following abbreviations are used for announcements: BI, business inventories; BOJ, Bank of Japan; BOP, balance of payments; CA, current account; CC, consumer credit; CCI, Consumer Confidence Index; CIPS, Chartered Institute of Purchasing and Supply (UK); COL, cost of living; CPI, Consumer Price Index; CS, construction spending; CU, capacity utilization; DG, durable goods orders; DR, discount rate; EC, employment costs; ECB, European Central Bank; EMU, European Monetary Union; FB, federal budget; FF, federal funds target; FI, factory inventories; FO, factory orders; FOMC, Federal Open Market Committee; FRB, Federal Reserve Bank; FS, factory shipments; GB; government budget; GDP, gross domestic product; GNP, gross national product; HC, housing completions; HCPI, Harmonized Consumer Price Index; HS, housing starts; IC, installment credit; IFO, Information and Forschung (Research) Institute (Germany); INSEE, French International Institute for Statistics and Economic Studies; IO, industrial orders; IP, industrial production; ISM, Institute for Supply Management Manufacturing Index; IUC, initial unemployment claims; LI, Index of Leading Indicators; M, imports; M1; M2; M3; M4; MI, Michigan Sentiment; MO, manufacturing orders; MOUT, manufacturing output;

Asset	Moment/Measure	Sample	Data frequency	Macro announcement(s)
USD/DEM USD/JPY USD/CHF USD/GBP USD/CAD	Conditional volatility/ Implied volatility	August 1986– April 1989	Daily	TB
USD/GBP	Conditional mean and volatility/ GARCH-M	April 1989– July 1989	Tick-by-tick	U.S.: TB U.K.: Interest rate
USD/DEM	Conditional mean and volatility/ Absolute returns and SDs	November 1988– November 1991	5-minute	CPI, DG, NFP, GNP, HS, MTB, LI, PPI, RS, IP, CU, BI, CS, FI, NAPM, NHS, PI, FB, IC
USD/JPY	Conditional volatility/ GARCH(1,1)	December 1983– February 1989	Daily (quotes from four markets)	TB

MPC, Monetary Policy Committee (UK); MPI, Import Price Index; MTB, merchandise trade balance; NAHB, National Association of Home Builders Housing Index; NAPM, National Association of Purchasing Managers Survey; NFP, nonfarm payroll employment; NHS, new home sales; NPROD, nonfarm productivity; OPEC, Organization of Petroleum Exporting Countries; PCE, personal consumption expenditures; PHI, Philadelphia Fed Index; PI, personal income; PMI, Chicago Purchasing Managers' Index; PPI, Producer Price Index; PSNCR, public sector net cash requirement; RA, reserve assets; RE, real earnings; RPIX, Retail Prices Index excluding mortgage interest payments; RPMI, Reuters Purchasing Managers' Index; RR, repo rate; RS, retail sales, TANKAN, quarterly poll of business confidence reported by the Bank of Japan; TB, trade balance; UR, unemployment rate; VR, new vehicle registration; VS, vehicle sales; WPI, Wholesale Price Index; WSS, wholesale sales; WSI, wholesale inventories; WST, wholesale turnover; X, exports; XPI, Export Price Index; ZEW, Centre for European Economic Research. The following abbreviations are used for currencies: ARS, Argentinean nuevo peso; AUD, Australian dollar; CAD, Canadian dollar; CHF, Swiss franc; CZK, Czech crown (koruna); DEM, Deutsche Mark; ECU, European Currency Unit; EUR, euro; FRF, French franc; GBP, British pound sterling; GRD, Greek drachma; HUF, Hungarian forint; IDR, Indonesian rupiah; ITL, Italian lira; JPY, Japanese yen; KRW, Korean won ; MXN, Mexican new peso; PLN, Polish new zloty; THB, Thai baht; TRY, Turkish new lira; TWD, Taiwan dollar; USD, U.S. dollar; ZAR, South African rand.

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APPENDIX, cont'd

Summary of the Literature on Estimating Announcement Effects on the Conditional Volatility and Jumps of Exchange Rate Returns

Reference	Abstract/Description*
Ederington and Lee (1994)	The paper examines the impact of major U.S. macroeconomic announcements on the Dollar/Yen exchange rate. We find that these announcements are responsible for most intraday and day-of-the-week volatility patterns in this market and we identify the most important announcements. The initial reaction to a major 8:30 announcement begins around 8:30:10 and lasts until about 8:30:50. A partial price correction is normally observed between 8:31 and 8:32. Price movements after 8:32 are basically independent of those observed earlier although volatility continues to be higher than normal until about 8:55.
Sultan (1994)	The objective of this study is to analyze the effects of trade deficit announcements on the joint distribution of the spot and futures price changes. In addition, this study examines whether or not trade deficits and trade surpluses have asymmetric effects on the currency price changes and volatility.
Johnson and Schneeweis (1994)	This study provides an examination of the effect of public news on inter-day exchange-rate return volatility. Unlike previous studies, the impacts of both U.S. and foreign macroeconomic news announcements are examined in the currency futures market for the Japanese yen, British pound, and Deutsche mark. Diffusion and jump-diffusion process models are developed which contain parameters conditional on the release of news…The results reveal that conditional variance diffusion and jump-diffusion process models dominate the equivalent non-conditional models…Thus, this study provides evidence that the currency return generating process is not characterized by a simple diffusion process over trading and non-trading periods. Further, the release of U.S. and foreign macroeconomic news has been shown to provide additional understanding of the currency return process over and above that of more complex models such as a jump-diffusion process.
Ederington and Lee (1995)	We examine how prices in interest rate and foreign exchange futures markets adjust to the new information contained in scheduled macroeconomic news releases in the very short run. Using 10-second returns and tick-by-tick data, we find that prices adjust in a series of numerous small, but rapid, price changes that begin within 10 seconds of the news release and are basically completed within 40 seconds of the release. There is some evidence that prices overreact in the first 40 seconds but that this is corrected in the second or third minute after the release. While volatility tends to be higher than normal just before the news release, there is no evidence of information leakage.
Leng (1996)	This article presents how the dollar/mark and dollar/yen exchange rates react to the anticipated U.S. monthly macroeconomic announcements…First, for both currencies, the 7:30-7:35 interval, immediately after the major announcements, not only has the largest average AAC, NP, and PR and is the only trading interval with positive average FAC. The impact of seven major announcements on these four price statistics lasts for at least an hour. On the other hand, the impact of the other 11 minor announcements is rather short lived.
Payne (1996)	This paper examines two aspects of spot FX [foreign exchange] volatility. Using intra-daily quotation data on the Deutsche Mark/dollar we simultaneously estimate the deterministic intra-daily seasonal pattern inherent in volatility and the effects of U.S. macroeconomic announcements. The empirical specification and estimation technique is based on the stochastic volatility methodology contained in Harvey, Ruiz, and Shephard (1994). Results conform with previous work, in that "news" effects are strong and persistent, being felt for over one hour after the initial release time. Inclusion of an explicit seasonal is shown to be essential for the accurate estimation of other volatility components. Further estimations allow us to examine which particular pieces of U.S. data move the markets. These results show that the most important statistics are those associated with the Employment and Mercantile Trade reports.

See NOTE on pp. 386-87.

Asset	Moment/Measure	Sample	Data frequency	Macro announcement(s)
USD/JPY	Conditional mean and volatility/ Absolute returns and SDs	November 1988– June 1993	10-second 5-minute 30-minute	CPI, DG, NFP, GNP, HS, LI, MTB, PPI, RS, IP, CU, BI, CS, FI, NAPM, NHS, PI, MPI, XPI, FB, IC
USD/CHF USD/CAD USD/DEM USD/JPY USD/GBP	Conditional mean and volatility/ Bivariate GARCH	February 1980– April 1989	Daily	TB
USD/JPY USD/GBP USD/DEM	Conditional volatility and jumps/ Jump-diffusion process and GARCH	January 1988– December 1990	Daily	U.S.: MTB, IP, CPI, Money supply U.K.: MTB, IP, CPI, Money supply Germany: MTB, IP, CPI, Money supply Japan: MTB, IP, CPI, Money supply
USD/DEM	Conditional mean and volatility/ Variance	November 1988– October 1992	Tick-by-tick 10-second	CPI, DG, NFP, UR, GNP, HS, LI, MTB, PPI, RS, IP, CU, BI, CS, FI, NAPM, NHS, PI, PCE, FB
USD/JPY USD/DEM	Conditional mean and volatility/ Absolute returns and period price ranges	November 1988– December 1993	5-minute	CPI, DG, NFP, GNP, HS, MTB, LI, PPI, RS, IP, CU, BI, CS, FI, NAPM, NHS, PI, FB, IC
USD/DEM	Conditional volatility/ Squared returns and ARCH	October 1992– September 1993	5-minute	UR, PPI, CPI, RS, CCI, LI, DG, IP, CU, NAPM, MTB

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APPENDIX, cont'd

Summary of the Literature on Estimating Announcement Effects on the Conditional Volatility and Jumps of Exchange Rate Returns

Reference	Abstract/Description*
DeGennaro and Shrieves (1997)	This paper estimates the impact of market activity and news on the volatility of returns in the exchange market for Japanese Yen and U.S. dollars. We examine the effects of news on volatility before, during and after news arrival; using three categories of news…Results indicate that both components of market activity, as well as news release, affect volatility levels. We conclude that both private information and news effects are important determinants of exchange rate volatility. Our finding that unexpected quote arrival positively impacts foreign exchange rate volatility is consistent with the interpretation that unexpected quote arrival serves as a measure of informed trading. Corroborating this interpretation is regression analysis, which indicates that spreads increase in the surprise component of the quote arrival rate, but not in the expected component. The estimated impact of a unit increase in unexpected quote arrival and the range of values observed for this variable imply an important volatility conditioning role for informed trading.
Andersen and Bollerslev (1998)	This paper provides a detailed characterization of the volatility in the deutsche mark-dollar foreign exchange market using an annual sample of five-minute returns. The approach captures the intraday activity patterns, the macroeconomic announcement, and the volatility persistence (ARCH) known from daily returns. The different features are separately quantified and shown to account for a substantial fraction of return variability, both at the intraday and daily level. The implications of the results for the interpretation of the fundamental "driving forces" behind the volatility process are also discussed.
Eddelbüttel and McCurdy (1998)	This paper investigates the impact of the frequency of general and currency-specific news headlines on deseasonalized intraday DEM-USD exchange rate changes. We find a significant relationship between volatility and the frequency of news. In particular, more news is associated with an increase in volatility. The result that spot exchange rates are more volatile during periods for which there is a lot of economic news accords with market participants' explanations for observed volatility clustering.
Joines, Kendall, and Kretzmer (1998)	Returns on a wide variety of assets are more volatile during trading hours than during nontrading hours. French and Roll (1986) suggest that this phenomenon may be due to (1) the concentrated release of public information, (2) the incorporation of private information into asset prices, or (3) the presence of trading noise. This paper uses returns on currency futures contracts traded on the Chicago Mercantile Exchange to study the importance of these three explanations. Unlike previous studies of currency futures, this paper finds evidence that either private information or trading noise is required to explain the observed pattern of return variance.
Kim (1998)	This paper examines the effects of scheduled Australian and U.S. macroeconomic announcements on daily USD/AUD exchange rate changes. EGARCH(1,1) models are used to investigate news effects on the conditional mean and volatility of the changes over various time horizons encompassing the announcements… The conditional volatility was higher in response to the Australian current account deficit and inflation news, while the retail sales news lowered it. The U.S. announcements, in general, had little effect during the U.S. market trading.

See NOTE on pp. 386-87.

Asset	Moment/Measure	Sample	Data frequency	Macro announcement(s)
USD/JPY	Conditional volatility/ SDs and GARCH	October 1992– September 1993	10-minute	U.S.: GDP, UR, DG, PPI, LI, RS, TB, CPI, Treasury bill auction, stimulus and tax news, interest rates Japan: GNP, IP, WPI, UR, LI, RS, TB, HS, CPI, stimulus and tax news, interest rates Bilateral: U.S. and Japan trade negotiations
USD/DEM	Conditional volatility/ GARCH	October 1992– September 1993	5-minute	U.S.: NFP, UR, DG, GDP, MTB, GDP, PPI, RS, LI, HS, FO, IUC, Treasury Report, CCI, CPI, CS, VS, BI, HC, MPI, MI, CA, CU, NPROD, M2, PI, RE, RA, HS, FOMC, Capital Spending Survey, NAPM, CC, WSS Germany: GDP, M3, Bundesbank meeting, WST, RS, CPI, IO, PPI, WPI, CA, Business insolvencies, NFP, MPI Japan: GNP
USD/DEM	Conditional volatility/ GARCH	October 1992– September 1993	Tick-by-tick	Headlines including keywords, such as "U.S." "DOLLAR," "FED"
USD/CAD USD/GBP USD/CHF USD/DEM USD/JPY	Conditional volatility/ SDs	January 1978– December 1987	Daily	None
USD/AUD	Conditional mean and volatility/ EGARCH	February 1985– April 1995	Daily	U.S.: CPI, GDP, UR, RS, TB Australia: CPI, GDP, UR, RS, CA

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APPENDIX, cont'd

Summary of the Literature on Estimating Announcement Effects on the Conditional Volatility and Jumps of Exchange Rate Returns

Reference	Abstract/Description*
Kim (1999)	This paper investigates the role of Australian macro-economic announcement news on five major Australian dollar (AUD) exchange rates…Current account deficit, CPI and unemployment news announcements significantly raised the conditional volatility of the changes of the AUD on the days of their announcements, except for the BP [British pound]/AUD for the CPI news, and there is some evidence of retail sales news reducing it. In general, the evidence is consistent with a view that a release of new information creates uncertainty in the markets due to a lack of consensus on the effects of the particular news announcement and the necessary course of action.
Han, Kling, and Sell (1999)	Using standard deviations and numbers of price changes calculated from tick data for currency futures, this study finds strong day-of-the week effects for the Deutsche mark and Japanese yen, mild effects for the British pound, and no effects for the Canadian dollar after controlling for scheduled macroeconomic announcements and days to contract expiration. The day-of-the-week effects are found to be caused either by Mondays' low volatility, or by Thursdays' or Fridays' high volatility. This result suggests that the day-of-the-week effects in the currency futures are not driven by the announcements of macroeconomic indicators as proposed in previous studies, but rather by other factors, such as private information-based trading or by market microstructure. This study also finds that the announcements are processed equally across the days of the week for all four currency futures.
Melvin and Yin (2000)	The mixture of distributions model motivates the role of public information arrival in foreign exchange market dynamics. Public information arrival is measured using Reuters Money Market Headline News. The exchange rates are high-frequency mark/dollar and yen/dollar quotes. Estimation results suggest that higher than normal public information brings more than the normal quoting activity and volatility. The results have implications for the debate over regulation of the foreign exchange market. Foreign exchange activity is not largely self generating. Trading is likely providing the function it is meant to provide—adjusting prices and quantities to achieve an efficient allocation of resources.
Cai et al. (2001)	The dramatic yen/dollar volatility of 1998 has been popularly ascribed to order flow driven by changing tastes for risk and hedge-fund herding on unwinding yen "carry trade" positions rather than fundamentals. High-frequency evidence of shifting fundamentals is provided by a comprehensive list of macroeconomic announcements. News is found to have significant effects on volatility, but order flow may play a more important role. Since portfolio shifts are revealed to the market through trading, the results are consistent with order flow playing a significant role in the revelation of private information and associated exchange rate shifts.

See NOTE on pp. 386-87.

Asset	Moment/Measure	Sample	Data frequency	Macro announcement(s)
USD/AUD DEM/AUD JPY/AUD GBP/AUD CHF/AUD	Conditional mean and volatility/ EGARCH	January 1985–April 1995	Daily	Australia: CA, CPI, GDP, UR, RS
USD/GBP USD/CAD USD/DEM USD/JPY	Conditional volatility/ SDs and number of price changes	January 1990–December 1997	Tick-by-tick	CPI, DG, NFP, GDP, MTB, PPI, RS, BI, CU, CS, FB, HS, IP, LI, PI, NAPM, NHS, FO, FI, FS
USD/DEM USD/JPY	Conditional volatility/ GARCH	December 1993–April 1995	Hourly	International: Reuters Money Market Headline News headlines related to U.S., Germany, or Japan
USD/JPY	Conditional volatility and order flow/ Absolute returns	January 1998–December 1998	5-minute	U.S.: NFP, CPI, GDP, PPI, MTB, IC, PI, CS, FOMC, NAPM, IP, IUC, HC, FB, DG, LI, M2, CCI, RS, BI, CA, VS, RE, EC, RA Japan: TANKAN, UR, VS, Consumer sentiment index, CA, vehicle exports, RA, HS, GDP, MTB, Tokyo new condo sales, vehicle imports, M2, LI, Tokyo department store sales, bank lending, bankruptcies, average lending rate, WPI, nationwide department store sales, corporate service price, Tokyo office vacancy rate, steel production, machine tool orders, household spending, IP, machinery orders, international securities investment, crude oil imports, CPI, electricity usage, vehicle production

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Summary of the Literature on Estimating Announcement Effects on the Conditional Volatility and Jumps of Exchange Rate Returns

Reference	Abstract/Description*
Ederington and Lee (2001)	We explore the determinants of intraday volatility in interest-rate and foreign-exchange markets, focusing on the importance and interaction of three types of information in predicting intraday volatility: (a) knowledge of recent past volatilities (i.e., ARCH or Autoregressive Conditional Heteroskedasticity effects); (b) prior knowledge of when major scheduled macroeconomic announcements, such as the employment report or Producer Price Index, will be released; and (c) knowledge of seasonality patterns. We find that all three information sets have significant incremental predictive power, but macroeconomic announcements are the most important determinants of periods of very high intraday volatility (particularly in the interest-rate markets). We show that because the three information sets are not independent, it is necessary to simultaneously consider all three to accurately measure intraday volatility patterns. For instance, we find that most of the previously documented time-of-day and day-of-the-week volatility patterns in these markets are due to the tendency for macroeconomic announcements to occur on particular days and at particular times. Indeed, the familiar U-shape completely disappears in the foreign-exchange market. We also find that estimates of ARCH effects are considerably altered when we account for announcement effects and return periodicity; specifically, estimates of volatility persistence are sharply reduced. Separately, our results show that high volatility persists longer after shocks due to unscheduled announcements than after equivalent shocks due to scheduled announcements, indicating that market participants digest information much more quickly if they are prepared to receive it. However, contrary to results from equity markets, we find no evidence of a meaningful difference in volatility persistence after positive or negative price shocks.
Evans (2002)	I examine the sources of exchange rate dynamics by focusing on the information structure of FX trading. This structure permits the existence of an equilibrium distribution of transaction prices at a point in time. I develop and estimate a model of the price distribution using data from the Deutsche mark/dollar market that produces two striking results: (1) Much of the short-term volatility in exchange rates comes from sampling the heterogeneous trading decisions of dealers in a distribution that, under normal market conditions, changes comparatively slowly; (2) public news is rarely the predominant source of exchange rate movements over any horizon.
Fornari et al. (2002)	This paper analyzes the impact of scheduled and unscheduled news on several Italian financial variables, paying particular attention to the effect on the conditional volatility of these variables. The impact of political and economic news items is assessed within a trivariate GARCH scheme. Results show that: (i) news affects both the first and the second moment of the daily changes in the analyzed variables; (ii) there is a significant regime shift of the unconditional variance of the analyzed variables across the three different governments in charge over the analyzed sample; (iii) the conditional variances display a significant (albeit rather small) seasonal daily pattern; (iv) contrary to the conventional view, the impact of news on the conditional variance is more pronounced for exchange rates than for Italian long-term interest rates.
Fair (2003)	Tick data and newswire searches are used to find events that led to large and rapid price changes in a stock future, a bond future, and three exchange rate futures. Knowledge of these events may be useful in future work. They have the advantages that they are truly surprises and that the sign of their effect on each financial instrument is known. The events are used in this study to analyze the effects of three types of events (monetary, price, and real) on the five instruments.

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Asset	Moment/Measure	Sample	Data frequency	Macro announcement(s)
USD/DEM	Conditional volatility/ ARCH	July 1989– May 1993	10-minute	CPI, DG, GNP, HS, LI, UR, MTB, FB, PPI, RS, IP, CU, CS, FI, NAPM, NHS, IP
USD/DEM	Conditional volatility and order flow/ Average number of interdealer transactions per minute, SDs	May 1996– August 1996	5-minute	International: Common knowledge and non-common knowledge news
USD/DEM ITL/DEM	Conditional volatility/ Trivariate GARCH	March 1994– November 1996	Daily	Italy: CPI, PPI, RR, 3-month bill auction rate, unscheduled news related to: public finance, prices, institutional conflicts, political conflicts, electoral results, easing of political tensions during the Dini government, Dini government, political debate beginning in 1996, EMU events directly related to Italy, EMU and the international environment, debate on judiciary procedures, the lira's Black Friday, Justice Minister Mancuso's actions and his parliamentary removal
USD/JPY USD/GBP USD/DEM-EUR	Conditional volatility and jumps/ Largest price changes and SDs	April 1982– March 2000	1-minute 5-minute	U.S.: NFP, UR, CPI, PPI, TB, RS, DG, HS, EC, GDP, IP, M1, NAPM, FF Japan: GDP, PCE, DR Germany: CPI, DR

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Summary of the Literature on Estimating Announcement Effects on the Conditional Volatility and Jumps of Exchange Rate Returns

Reference	Abstract/Description*
Andersen et al. (2003)	Announcement surprises produce conditional mean jumps; hence high-frequency exchange-rate dynamics are linked to fundamentals. The details of the linkage are intriguing and include announcement timing and sign effects. The sign effect refers to the fact that the market reacts to news in an asymmetric fashion: bad news has greater impact than good news.
Chang and Taylor (2003)	This paper investigates the link between information arrivals and intraday DEM/$ volatility. Information arrivals are measured by the numbers of news items that appeared in the Reuters News Service. We separate news stories into different categories and find that total headline news counts, U.S. and German macroeconomic news and German Bundesbank monetary policy news all have a significant impact on intraday DEM/$ volatility. The quantitative effects of the total news counts are less than those of the German Bundesbank monetary policy news and U.S. macroeconomic news. News related to the U.S. Federal Reserve appears to have little impact because of the Federal Reserve's steady monetary policy during the sample period. The conclusions are obtained from ARCH models that incorporate intraday seasonal volatility terms.
Han (2004)	Using new datasets of high frequency Dollar-Euro foreign exchange rates, surveyed expectations and actual realizations of macroeconomic indicators in the U.S. and the EMU, this paper characterizes a new type of the high frequency Dollar-Euro foreign exchange rate data after 1999 when the Euro currency was first introduced in foreign exchange markets. The FIGARCH model is found to be the preferred specification for the Dollar-Euro returns data, with similar values of the long memory volatility parameter across different frequencies, which is indicative of returns being generated by a self similar process. This paper also examines whether the Euro currency reacts to macroeconomic shocks in different ways depending on whether the shocks come from the U.S. or the EMU region and whether the shocks are positive or negative. By quantifying the duration of the intraday impacts of the macroeconomic shocks on the high frequency Dollar-Euro returns, this paper finds that the macroeconomic shocks of the U.S. and the EMU are found to have statistically significant impacts on both the conditional mean and the conditional variance but their impacts appear to be asymmetric depending on the regions (U.S. and EMU area) and the signs (positive and negative) of the shocks.
Kim, McKenzie, and Faff (2004)	We investigate the impact of scheduled government announcements relating to six different macroeconomic variables on the risk and return of three major U.S. financial markets. Our results suggest that these markets do not respond in any meaningful way, to the act of releasing information by the government. Rather, it is the "news" content of these announcements which cause[s] the market to react. For the three markets tested, unexpected balance of trade news was found to have the greatest impact on the mean return in the foreign exchange market. In the bond market, news related to the internal economy was found to be important. For the U.S. stock market, consumer and producer price information was found to be important. Finally, financial market volatility was found to have increased in response to some classes of announcement and fallen for others. In part, this result can be explained by differential "policy feedback" effects.
Kopecký (2004)	By using high frequency exchange rate data I examine the reaction of the Czech Crown/USD spot exchange rate to public macroeconomic announcements emanating from the U.S. and the Czech Republic. I directly test the efficient market hypothesis. By using data spaced at 5-minute intervals I identify significant impacts of the news on the exchange rate and its volatility, and test for the presence of announcement specific effects. Analysis of the volatility yields a spike in the ten minutes following the Czech announcements… The volatility of CZK/USD returns does not increase following the U.S. announcements.

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Asset	Moment/Measure	Sample	Data frequency	Macro announcement(s)
USD/CHF USD/DEM USD/EUR USD/JPY USD/GBP	Conditional mean and jumps/ Absolute returns	January 1992–December 1998	5-minute	U.S.: GDP, NFP, RS, IP, CU, PI, CC, PCE, NHS, DG, CS, FO, BI, GB, TB, PPI, CPI, CCI, NAPM, HS, LI, FF, IUC, M1, M2, M3 Germany: NFP, RS, IP, MOUT, MO, TB, CA, CPI, PPI, WPI, MPI, M3
USD/DEM	Conditional mean and volatility/ Variance, absolute returns, and GARCH	October 1992–September 1993	5-minute 10-minute 15-mintue 30-minute Hourly	U.S.: IUC, FB, BI, CS, CPI, DG, FO, GDP, HS, IP, CU, UR, LI, PI, PPI, PMI, RS, NHS, TB, FF, RA Germany: WST, IO, CPI, PPI, WPI, MPI, CA, UR, TB, RS, COL, engineering orders, M3, GDP, GNP, business insolvencies, RR, DR
USD/EUR	Conditional mean and volatility/ FIGARCH	January 1999–December 2002	15-minute	U.S.: UR, PPI, RS, CPI, IP, GDP Europe: UR, PPI, RS, CCI, GDP, IP, CPI
USD/DEM USD/JPY	Conditional mean and volatility/ GARCH	January 1986–December 1998	Daily	TB, GDP, UR, RS, CPI, PPI
USD/CZK	Conditional mean and volatility/ SDs of log returns	January 1997–December 2002	5-minute	U.S.: CPI, PPI, IP, MTB, UR, DG Czech Republic: PPI, CPI, IP, TB

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Summary of the Literature on Estimating Announcement Effects on the Conditional Volatility and Jumps of Exchange Rate Returns

Reference	Abstract/Description*
Sager and Taylor (2004)	We examine the evidence regarding systematic patterns in the euro-dollar foreign exchange market on days versus other days. We examine 5-minute data in a nonlinear framework allowing for switching between a high-volatility, informed-trading state and a low-volatility, liquidity-trading state. We find strong evidence that the GC policy announcements contain significant news content. Although there is some evidence of positioning in the hour prior to the announcement, this probably reflects dealers minimizing their exposure rather than evidence of information leakage.
Laakkonen (2004)	This study investigates the impact of new information on the volatility of exchange rates. The impact of scheduled U.S. and European macroeconomic news on the volatility of USD/EUR 5-minute returns was tested by using the Flexible Fourier Form method. The results were consistent with earlier studies. Macroeconomic news increased volatility significantly, and news on the United States was the most important. The much-tested hypothesis of bad news having a greater impact on volatility was re-confirmed in this study. The announcements were also divided into two categories, the first containing the news that gave conflicting information on the state of the economy (bad and good news at the same time) and the other containing the news that was consistent (where either good or bad news was announced). Conflicting news was found to increase volatility significantly more than consistent news. The impact of "no-surprise" news was also tested. Even news the forecast of which was equal to an announcement seemed to increase volatility.
Bauwens, Ben Omrane, and Giot (2005)	We study the impact of nine categories of scheduled and unscheduled news announcements on the euro/dollar return volatility. We highlight and analyze the pre-announcement, contemporaneous and post-announcement reactions. Using high-frequency intraday data and within the framework of ARCH-type models, we show that volatility increases in the pre-announcement periods, particularly before scheduled events. Market activity also significantly impacts return volatility as expected by the theoretical literature on the order flow.

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Asset	Moment/Measure	Sample	Data frequency	Macro announcement(s)
USD/EUR	Conditional volatility/ GARCH	January 2002– December 2003	5-minute	Europe: ECB monthly meeting
USD/EUR	Conditional volatility/ SDs and GARCH	October 2003– January 2004	5-minute	U.S.: RS, Hourly earnings, weekly hours, building permits, BI, CU, NFP, PMI, CS, CCI, CC, CPI, CA, VS, DG, EC, FO, FOMC, GDP, HS, IP, IUC, LI, PI, RS, TB, MI, UR, WSI, MPI, ISM, NHS, NAHB, PCE, NPROD, PHI, PPI, VS, GB Europe: Production in construction, RPMI, business climate, RA, CPI, GDP, IP, EC, VR, RS, TB, UR, CCI, M3, interest rates, CA, GB, PPI, IO, LI, ZEW survey France: RPMI, CCI, CPI, PCE, CA, GDP, HS, IP, MOUT, NFP, PPI, UR, GB, TB, wages, VR Germany: RPMI, GB, construction investment, construction orders, CPI, CA, domestic demand, NFP, X, equipment investment, FO, GDP, IP, PPI, RS, UR, WPI, MPI, M, CCI, Ifo Business Climate Survey, VR, PCE, TB, ZEW survey
USD/EUR	Conditional volatility and order flow/ Variance and EGARCH	May 2001– November 2001	5-minute	U.S.: NFP, NAPM, WSS, GDP, PPI, RS, HS, CCI, CPI, CS, VS, BI, HC, MPI, CA, NPROD, PI, RE, NHS, speeches of senior officials of the government and of public agencies, interest rate report Europe: Unspecified macroeconomic figures, speeches of senior officials of the government and of public agencies, interest rate report International: forecasts made by economic institutes, declarations of OPEC members, rumors of central bank interventions, extraordinary events

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Summary of the Literature on Estimating Announcement Effects on the Conditional Volatility and Jumps of Exchange Rate Returns

Reference	Abstract/Description*
Jansen and De Haan (2005)	This paper studies the reaction of the conditional mean and volatility of the euro–dollar exchange rate to statements by European Central Bank and national central bank officials. We focus on comments on monetary policy and the external value of the euro. We find that the Bundesbank has dominated the news coverage. We conclude that ECB statements have mainly influenced conditional volatility. In some cases there are effects of statements on the conditional mean of the euro–dollar exchange rates. Efforts to talk up the euro have generally not been successful. There is also evidence of asymmetric reactions to news.
Evans and Lyons (2005)	News arrivals induce subsequent changes in trading in all of the major end-user segments. These induced changes remain significant for days. Induced trades also have persistent effects on prices. Currency markets are not responding to news instantaneously.
Carlson and Lo (2006)	A surprise announcement of an increase in German interest rates coupled with concurrent transactions data enables us to study in detail dealers' reactions. The patterns observed are consistent with dealers' practice to book targeted profits immediately if possible in the face of uncertainty. Evidence also shows that the speculative activity by traders in initial reaction to the news destabilized the market for the next 2 hours.
Dominguez and Panthaki (2006)	This paper examines whether the traditional sets of macro surprises, that most of the literature considers, are the only sorts of news that can explain exchange-rate movements. We examine the intra-daily influence of a broad set of news reports, including variables which are not typically considered "fundamentals" in the context of standard models of exchange-rate determination, and ask whether they too help predict exchange-rate behavior. We also examine whether "news" not only impacts exchange rates directly, but also influences exchange rates via order flow (signed trade volume). Our results indicate that along with the standard fundamentals, both non-fundamental news and order-flow matter, suggesting that future models of exchange-rate determination ought to include all three types of explanatory variables.
Savaser (2006)	I find that price-contingent orders can enhance our ability to explain post-release exchange-rate returns by half. Furthermore, the estimated effect of orders is orthogonal to the news surprises.
Pearce and Solakoglu (2007)	This paper examines the relationship between macroeconomic news and the dollar-Mark and dollar-Yen exchange rates…We examine the linearity and symmetry of the responses to news and also allow the effects of the news announcements to vary across states of the economy. We find that news indicating a stronger U.S. economy causes an appreciation of the U.S. dollar, that the responses are essentially complete within 5 min, and that measuring the responses over 6-h intervals eliminates the statistical significance of the news. The effects of news appear linear and symmetric but there is some evidence that the effects depend on the state of the economy.
Andersen et al. (2007)	We characterize the response of U.S., German and British stock, bond and foreign exchange markets to real-time U.S. macroeconomic news…[N]ews produces conditional mean jumps; hence high-frequency stock, bond and exchange rate dynamics are linked to fundamentals…when conditioning on the state of the economy, the equity and foreign exchange markets appear equally responsive…[W]e also document important contemporaneous links across all markets and countries, even after controlling for the effects of macroeconomic news.

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Asset	Moment/Measure	Sample	Data frequency	Macro announcement(s)
USD/EUR	Conditional mean and volatility/ EGARCH	January 1999– May 2002	Daily	Europe: headlines from Bloomberg News Service with keywords "ECB" or names of ECB officials, as well as keywords for national central banks and their officials
USD/EUR	Conditional mean, conditional volatility, and order flow/ Variance	April 1993– June 1999	Daily	U.S.: BI, CU, IUC, CCI, CS, CPI, CC, DG, FO, FF, GDP, RE, HS, IP, LI, M1, M2, M3, NAPM, NHS, NFP, PCE, PI, PPI, RS, GB, TB, UR Germany: GDP, NFP, RS, IP, MOUT, MO, TB, CA, COL, WPI, PPI, MPI, M3
USD/DEM	Conditional mean and order flow/ Absolute returns	October 1997	Tick-by-tick	Bundesbank interest rate hike
USD/EUR USD/GBP	Conditional mean, conditional volatility, and order flow/ GARCH	October 1999– July 2000	20-minute	U.S.: PPI, CPI, IP, M3, TB, UR, NFP, RS U.K.: RPIX, RS, TB, M4, PPI, IP, UR, CA Europe: PPI, CPI, IP, M3, TB, UR
USD/GBP	Conditional mean, conditional volatility, order flow, and jumps/ Absolute returns	September 1999– April 2000 and June 2001– September 2002	5-minute	GDP, NFP, RS, DG, BI, TB, PPI, CPI, HS, LI, PCE, PI, IUC
USD/DEM USD/JPY	Conditional mean and volatility/ Variance	December 1986– December 1996	5-minute	CPI, PPI, M2, TB, UR, IP, CCI, DG, NAPM, RS, NFP
USD/GBP USD/JPY USD/DEM-EUR	Conditional mean and volatility/ SDs	January 1992– December 2002	5-minute	GDP, NFP, RS, IP, CU, PI, CC, NHS, PCE, DG, FO, CS, BI, FB, TB, PPI, CPI, CCI, NAPM, HS, LI, FF, IUC

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Summary of the Literature on Estimating Announcement Effects on the Conditional Volatility and Jumps of Exchange Rate Returns

Reference	Abstract/Description*
Chaboud, Chernenko, and Wright (2008)	This article introduces a new high-frequency data set that includes global trading volume and prices over five years in the spot euro-dollar and dollar-yen currency pairs. Studying the effects of U.S. macroeconomic data releases, we show that spikes in trading volume tend to occur even when announcements are in line with market expectations, in sharp contrast to the price response. There is some evidence that the volume after announcements is negatively related to the ex ante dispersion of market expectations, contrary to the standard theoretical prediction. At very high frequency, we find evidence that much of the immediate jump in prices in reaction to an announcement occurs before the surge in volume.
Frömmel, Mende, and Menkhoff (2008)	This paper examines the roles of order flow (reflecting private information) and news (reflecting public information) in explaining exchange rate volatility. Analyzing four months of a bank's high frequency dollar/euro trading, three different kinds of order flow are used in addition to seasonal patterns in explaining volatility. We find that only larger sized order flows from financial customers and banks—indicating informed trading—contribute to explaining volatility, whereas flows from commercial customers do not. The result is robust when we control for news and other measures of market activity. This strengthens the view that exchange rate volatility reflects information processing.
Hayo and Neuenkirch (2009)	Canadian and U.S. price shocks and monetary policy news are less important than shocks relating to the real economy…Canadian central bank communication is more relevant than its U.S. counterpart, whereas in the case of macro news that originating from the United States dominates…[T]he impact of Canadian news reaches its maximum when the Canadian target rate departs from the Federal Funds target rate (2002-2004) and thereafter.
Melvin et al. (2009)	We find evidence for non-linear regime switching between a high-volatility, informed-trading state and a low-volatility, liquidity-trading state. MPC surprise announcements are shown significantly to affect the probability that the market enters and remains within the informed trading regime.
Berger, Chaboud, and Hjalmarsson (2009)	We propose a new empirical specification of volatility that links volatility to the information flow, measured as the order flow in the market, and to the price sensitivity to that information. The time-varying market sensitivity to information is estimated from high-frequency data, and movements in volatility can therefore be directly related to movements in order flow and market sensitivity. Empirically, the model explains a large share of the long-run variation in volatility. Importantly, the time variation in the market's sensitivity to information is at least as relevant in explaining the persistence of volatility as the rate of information arrival itself. This may be evidence of a link between changes over time in the aggregate behavior of market participants and the time-series properties of realized volatility.
Cai, Joo, and Zhang (2009)	This paper utilizes a unique high-frequency database to measure how exchange rates in nine emerging markets react to macroeconomic news in the U.S. and domestic economies from 2000 to 2006. We find that major U.S. macroeconomic news have a strong impact on the returns and volatilities of emerging market exchange rates, but many domestic news do not. Emerging market currencies have become more sensitive to U.S. news in recent years. We also find that market sentiment could sway the impact of news on these currencies systematically, as good (bad) news seems to matter more when optimism (pessimism) prevails. Market uncertainty also interacts with macroeconomic news in a statistically significant way, but its role varies across currencies and news.

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Asset	Moment/Measure	Sample	Data frequency	Macro announcement(s)
USD/EUR USD/JPY	Conditional mean, conditional volatility, and jumps/ Variance of returns	January 1999–February 2004	1-minute	U.S.: GDP, PPI, NFP, TB, RS, FF
USD/EUR	Conditional volatility and order flow/ GARCH(1,1)	July 2001–November 2001	1-minute	U.S./Euro: GDP, NFP, CPI, RS, MI, CCI, Statements of central banks and other institutions on the whole economy
USD/CAD EUR/CAD	Conditional mean and volatility/ GARCH-M	January 1998–December 2006	Daily	U.S.: Federal Reserve Board of Governors' Statements, GDP, IP, TB, ISM, CCI, HS, NFP, UR, CPI, PPI, RS, FF Canada: Canadian Governing Council's statements, GDP, CU, CA, MTB, Ivey Purchasing Managers Index, HS, NFP, UR, RS, CPI, Industrial Product Price Index, Raw Materials Price Index, Central bank target interest rates
USD/GBP	Conditional mean and volatility/ GARCH	June 1997–October 2007	5-minute Daily	U.K.: MPC meeting
USD/EUR	Conditional volatility and order flow/ Integrated volatility	January 1999–December 2004	1-minute 5-minute	International: Order flow
USD/CZK USD/HUF USD/IDR USD/KRW USD/MXN USD/PLN USD/ZAR USD/THB USD/TRY	Conditional mean and volatility/ GARCH(1,1)	January 2000–December 2006	5-minute	U.S.: BI, GB, CA, CU, CCI, CC, CS, CPI, DG, FO, GDP, HS, imports, FF, IP, NAPM, LI, NHS, NFP, PCE, PI, PPI, RS, TB, IUC, WSS, Emerging markets (where available): GB, CA, CCI, CPI, X, fixed investment, GDP, M, interest rate, IP, money supply, PPI, RS, TB, IUC, WSS

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Summary of the Literature on Estimating Announcement Effects on the Conditional Volatility and Jumps of Exchange Rate Returns

Reference	Abstract/Description*
Chelley-Steeley and Tsorakidis (2009)	In January 2001 Greece joined the eurozone. The aim of this article is to examine whether an intention to join the eurozone had any impact on exchange rate volatility. We apply the Iterated Cumulative Sum of Squares (ICSS) algorithm of Inclan and Tiao (1994) to a set of Greek drachma exchange rate changes. We find evidence to suggest that the unconditional volatility of the drachma exchange rate against the dollar, British pound, yen, German mark and ECU/Euro was nonstationary, exhibiting a large number of volatility changes prior to European Monetary Union (EMU) membership. We then use a news archive service to identify the events that might have caused exchange rate volatility to shift. We find that devaluation of the drachma increased exchange rate volatility but ERM [European Exchange Rate Mechanism] membership and a commitment to joining the eurozone led to lower volatility. Our findings therefore suggest that a strong commitment to join the eurozone may be sufficient to reduce some exchange rate volatility which has implications for countries intending to join the eurozone in the future.
Harada and Watanabe (2009)	This paper studies the high frequency reaction of the Yen/Dollar exchange rate to announced macroeconomic information emanating from Japan and the U.S. We use data sampled at a five-minute frequency over a period from six years, from January 2001 to December 2006. We find that only announced surprises of the U.S. news produce impacts both on the conditional returns and its volatilities of the Yen/Dollar exchange rate and that those of Japanese news have much less impacts. The effects on the returns and those on the volatilities appear with delay.
Hashimoto and Ito (2009)	Using high-frequency transaction data of the actual trading platform, we examine market impact of Japanese macroeconomic statistics news within minutes of their announcements on the dollar/yen exchange rate. Macroeconomic statistics surprises that consistently have significant effect on dollar/yen returns include TANKAN (business condition survey conducted by Bank of Japan), GDP, industrial production, price indices and balance of payment. The announcement itself, in addition to the magnitude of the surprise, is found to increase the number of deals and price volatility immediately after the announcement. Most effects, when significant, take place within 30 min of statistics announcements.
Laakkonen and Lanne (2009)	We study whether the accuracy of news announcements matters for the impact of news on exchange rate volatility. We use high-frequency EUR/USD returns and releases of 20 U.S. macroeconomic indicators, and measure the precision of news in three different ways. When the precision is defined by the size of the first revision of the previous month's figure, we find that precise news increases volatility significantly more than imprecise news. Also, news on indicators that are in general more precise increase volatility more than news on typically imprecise indicators. Finally, we use real time data to measure the "true" precision of news and find that the size of the first revision of the previous month's figure is a reasonable signal of "true" precision.
Conrad and Lamla (2010)	We investigate the impact of the European Central Bank's monetary policy communication during the press conference held after the monthly Governing Council meeting on the EUR-USD exchange rate in high frequency. Based on the method of Content Analysis we construct communication indicators for the introductory statement and find that communication with respect to future price developments is most relevant. In response to statements about increasing risks to price stability the EUR appreciates on impact. To the contrary, communication about economic activity and monetary aggregates does not generate significant exchange rate reactions.

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Asset	Moment/Measure	Sample	Data frequency	Macro announcement(s)
USD/GRD GBP/GRD JPY/GRD DEM/GRD EUR/GRD	Conditional volatility/ Mean-centered cumulative sum of squares	January 1995– December 2000	Daily	Greece: The intention of Greece to join the euro zone
USD/JPY	Conditional mean and volatility/ SDs	January 2001– December 2006	5-minute	U.S.: GDP, CPI, IP, UR, FF Japan: GDP, CPI, IP, UR, Monetary policy of BOJ, TANKAN
USD/JPY	Conditional mean and volatility/ Sum of squared returns	January 2001– December 2005	1-minute 5-minute 15-minute 30-minute	Japan: CPI, UR, TANKAN, M2+, GDP, BOP, TB, PPI, IP, NFP, Diffusion Index, PCE, HS, RS
USD/EUR	Conditional volatility/ GARCH	January 1999– December 2004	5-minute	CU, NFP, PMI, CPI, CCI, DG, FO, GDP, HS, MPI, IP, IUC, ISM, LI, NHS, PPI, TB, MI, WSI, PHI
USD/EUR	Conditional mean and volatility/ Absolute returns, FIGARCH, and AR-FIGARCH	January 1999– October 2006	5-minute	Europe: ECB press releases

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APPENDIX, cont'd

Summary of the Literature on Estimating Announcement Effects on the Conditional Volatility and Jumps of Exchange Rate Returns

Reference	Abstract/Description*
Evans and Speight (2010)	The short-run reaction of Euro returns volatility to a wide range of macroeconomic announcements is investigated using five-minute returns for spot Euro-Dollar, Euro-Sterling and Euro-Yen exchange rates. The marginal impact of each individual macroeconomic announcement on volatility is isolated whilst controlling for the distinct intraday volatility pattern, calendar effects, and a latent, longer run volatility factor simultaneously. Macroeconomic news announcements from the U.S. are found to cause the vast majority of the statistically significant responses in volatility, with U.S. monetary policy and real activity announcements causing the largest reactions of volatility across the three rates. ECB interest rate decisions are also important for all three rates, whilst U.K. Industrial Production and Japanese GDP cause large responses for the Euro-Sterling and Euro-Yen rates, respectively. Additionally, forward looking indicators and regional economic surveys, the release timing of which is such that they are the first indicators of macroeconomic performance that traders observe for a particular month, are also found to play a significant role.
Lahaye, Laurent, and Neely (2010)	Nonfarm payroll and federal funds target announcements are the most important news across asset classes. Trade balance shocks are important for foreign exchange jumps. We relate the size, frequency and timing of jumps across asset classes to the likely sources of shocks and the relation of asset prices to fundamentals in the respective classes.

See NOTE on pp. 386-87.

Asset	Moment/Measure	Sample	Data frequency	Macro announcement(s)
USD/EUR GBP/EUR JPY/EUR	Conditional volatility/ GARCH	January 2002– July 2003	5-minute	U.S.: BI, Challenger layoffs, Chicago National Activity Index, PMI, CS, CCI, CC, CPI, CA, DG, NFP, UR, existing home sales, FO, FI, FOMC, GDP, HS, HC, housing permits, MPI, XPI, IP, CU, IUC, ISM, LI, MI, M2, NAHB, NHS, PI, PCE, PHI, PPI, NPROD, RS, TB, FB Europe: Business climate index, CCI, Business Confidence Index, Sentiment Index, CPI, CA, GDP, HCPI, IP, EC, M3, LI, PPI, PMI, RS, Services Index, Composite Index, TB, UR, ECB Germany: CA, COL, Capital account, NFP, GDP, Ifo business expectations, Ifo Manufacturing Survey, MPI, IP, MO, PMI, PPI, RS, Services Index, TB, UR, ZEW expectations France: Business climate, CPI, CA, GDP, PCE, Household Survey, IP, INSEE report, NFP, PPI, PMI, Services Index, TB, UR U.K.: CIPS Manufacturing Survey, CIPS Services Survey, CCI, CC, GDP, CA, Halifax House Price Index, IP, MO, M4, MPC, Nationwide house prices, PPI, PSNCR, RS, RPIX, TB, UR Japan: BOJ, Coincident Index, Construction orders, HS, CCI, CPI, Department store sales, FX Reserves, GDP, PI, IP, M2, RS, Shipments, Supermarket sales, TANKAN, TB, CA, Tertiary Index, UR
USD/EUR USD/GBP USD/JPY USD/CHF	Conditional mean and jumps/ GARCH, ARCH, and tobit-GARCH	January 1987– October 2004	5-minute	GDP, NFP, RS, IP, CU, CC, PI, PPI, CPI, DG, BI, CS, FO, PCE, NHS, TB, GB, Manufacturing Composite Index, HS, CCI, LI, FF

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