The Problem of Small Change in Early Argentina

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Many economies, during the early stages of monetary development, experienced what appear to be sporadic relative shortfalls of small denomination means of payment. These episodes have been broadly documented in the literature under the label of "shortages of small change." Sargent and Velde (2002), for example, review in great detail the evidence for Europe. Hanson (1979) provides an interesting survey of the evidence for the British colonies in North America. The purpose of this paper is to present evidence of similar events occurring in the early monetary history of Argentina.

The provision of small change in modern economies has become almost a nonissue.[1] In fiat money systems, the monetary authority controls the aggregate supply of monetary balances (of all denominations) and stands ready to exchange at par any denomination for an equivalent amount of any other denomination. It is, then, demand that determines the relative amounts of the different denominations that circulate in the economy. There are, of course, costs of providing the demanded amounts. Low-denomination coins tend to be relatively more costly to produce (at least, per unit of value). Yet, in general, governments in modern societies have considered these costs worthwhile.

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[1] Some high-inflation economies sometimes experience imbalances in their denomination structure, resulting in relative scarcity of small change. The monetary authority becomes reluctant to provide large quantities of low-denomination means of payment, which are relatively costly to produce and, due to high inflation, have a short useful lifespan.

Under a commodity money system it is also possible, in principle, to solve the problem of the relative supply of denominations. Basically, it requires issuing token coins that can be readily exchanged at par with the monetary authority for full-bodied coins.[2] This monetary arrangement is commonly known as the "standard formula" (see Sargent and Velde 2002, 5). For some time, the production of token coins represented a significant technological challenge. Since token coins are worth more than their intrinsic value, counterfeiting is a very profitable activity under the "standard formula" system. To avoid counterfeiting, the coins need to be fairly sophisticated, which increases their production cost. In the history of monetary systems, before the technology for production of coins was well developed, token coins were at best a very imperfect solution to the problem of small change.

When all coins are full-bodied coins, as was the case in Spanish America, the potential for mismatches on the demand and supply of denominations becomes more likely. The minting cost per value is normally lower for high-denomination coins. To the extent that coins circulate at par value (i.e., without discounts), there are incentives to mint only high-denomination coins. The historical record on the shortages of small change is the story of monetary authorities that struggled to sustain the proper mix of (full-bodied) high- and low-denomination coins.

Imperfect private solutions to the shortage problem are possible. For example, coins could be cut in portions to circumvent the indivisibility barrier. Also, small change can be offered at a premium. However, all these partial solutions bring forth the problems that arise when high- and low-denomination coins do not exchange in convenient and fixed ratios. These issues are most relevant in cases where the need for small change originates in domestic transactions. Systematic negotiation over discounts and measurement of effective coin weight, in this case, involve very low-value amounts, most likely not worth the trouble.

In Search of a Formal Definition

Providing a precise theoretical definition of a shortage of small change is no simple matter. Sargent and Velde (1999; 2002) consider a model where low-denomination money can be used for "small" and "large" transactions but high-denomination money can only be used for "large" transactions. They then define a "shortage" as a situation where the agents in the economy have to adjust their consumption pattern to their holdings of low-denomination money while, at the same time, holding an "excess" stock of the high-denomination

[2] A full-bodied coin is one that has metallic content worth as much as the coin's face value. Token coins are subsidiary coins that have lower intrinsic value than the value at which they can be exchanged with the monetary authority for full-bodied coins.

money. In other words, at the time of consumption decisions, the agent's holdings of low-denomination money act as a binding constraint (see Rolnick and Weber [2003] for a good summary).

Such a definition provides an interesting prediction that can potentially be confronted with the data. Basically, shortages of low-denomination money (as defined by Sargent and Velde) coincide with the persistent depreciation of low-denomination money relative to high-denomination money. The economic intuition for this result is simple. During shortages, low-denomination money is effectively more useful in transactions than high-denomination money. However, both monies are being voluntarily held by agents. For this situation to be an equilibrium, the low-denomination money has to be losing value with respect to the high-denomination money. In other words, while low-denomination money is more useful in transactions, its value is depreciating with time. The two effects offset each other in equilibrium and leave agents indifferent (at the margin) between holding low- and high-denomination money.[3]

Wallace (2003) strongly criticizes Sargent and Velde's definition. Wallace points out that the historical "shortages are reports by contemporaries concerning the difficulty of carrying out trade in the face of a sudden disappearance of some kinds of coins." However, in Sargent and Velde's model, agents are "freely choosing quantities while taking prices as given," and "nothing in the model looks like a shortage or a disruption of trade."

Wallace goes on to discuss what is, in his view, at the heart of the problem of small change. He starts by specifying the four desirable properties of a medium of exchange: portability, divisibility, durability, and recognizability; and he describes the history of coinage as "mainly about the technological difficulties of achieving a full-bodied coinage system that comes close to having those attributes." Formal economic analyses that explicitly model *all four* attributes of a medium of exchange and study the denomination structure of money are not readily available. Wallace (2003) provides an introduction to the formal treatment of these matters within the framework provided by the random matching models of money pioneered by Kiyotaki and Wright (1989).[4] He discusses how the indivisibility of coins, for example, could limit the set of transactions that agents undertake and how such limitations would clearly be perceived as shortages of small change by the agents in the model. These are interesting theoretical avenues that should eventually improve our understanding of the historical records.

The objective of this article is not to answer the theoretical questions that surround the issue of the appropriate provision of small change. Those

[3] Sargent and Velde (2002) do, in fact, find some supporting evidence for this hypothesis in their review of the historical record of many European countries. Unfortunately, as we will later see, such an evaluation is not possible with the data available for the colonial period in South America.

[4] See also Wallace and Zhou (1997) and Lee, Wallace, and Zhu (2005).

are difficult questions that involve theoretical concepts that are not yet well understood in the literature and require much further study. Here, instead, we will limit ourselves to reviewing the historical evidence that indicates that the availability of small change was a problem around the turn of the 18th century in the region that is today Argentina. As noted above, similar evidence is available for other parts of the world. However, in view of the intricate nature of the problem involved, it seems important to collect and analyze as much evidence as possible on this matter.

A Caveat

Before turning to the main subject of the article, a general clarification is in order. When reading the historical records, one has to be especially careful in differentiating the scarcity of means of payment from the general scarcity of resources prevailing in the area. The confusion between those two different phenomena was common at the time (for example, official resolutions would sometimes associate the general scarcity of resources with the inability to issue coins).[5] The territory of Argentina was relatively poor at the end of the 18th and the beginning of the 19th centuries and it relied on imports from Europe for many essential needs. Furthermore, the recurrent military conflicts at the beginning of the 1800s only contributed to further reducing the availability of economic resources. But the general stringency of resources is not the subject of this article. Instead, I intend to report evidence that suggests that the relative scarcity of certain means of payment, and in particular *fractional* money, constituted a problem in itself.

The remainder of the article is organized as follows. In the next section, I discuss the evidence from 1776 to the 1810 Revolution, that is, the latter part of the colonial period. Classic manifestations of the scarcity of small change appear in this period: attempts by the government to ban the export of fractional money; extreme difficulties in persuading the population to remint low-quality, low-denomination coins; and the development of imperfect substitutes for use in payment of small transactions are among the main ones being reviewed. In Section 2, I present evidence that suggests that the shortage of small change continued to be a problem during the 15 years after the revolution. I explain how the government struggled with the decision to issue copper coins for 10 years, and how it finally injected the coins in 1823 with great initial success. Also, in 1822 the first bank in the region was created and allowed to issue notes of moderate denominations after some initial reluctance. I discuss the genesis of that decision in some detail. To conclude, in Section 3, I provide

[5] Supple (1957, 244–45) reports that this kind of confusion was also common in 17th century England.

some further discussion of the economic issues highlighted by the evidence discussed.

1. EVIDENCE FROM THE PERIOD OF THE VICEROYALTY OF THE RIO DE LA PLATA

Spanish settlers were in the area that is today Argentina since the mid-1500s. For the early colonial period, the available evidence indicates that there was no widespread monetary exchange occurring in the region and that instead barter was the predominant way of exchange (Prebisch 1921, 193; Elía 1942a, 416). Since the beginning of colonization, Spain implemented a system of international trade restrictions in the colonies. The port of Buenos Aires could only trade with Spain and such trade was subject to heavy taxation. These restrictions significantly slowed down economic development for more than two hundred years. Only in 1776, with the creation of the Viceroyalty of the Rio de la Plata, the representative of the king, Viceroy Ceballos, declared free trade in the port of Buenos Aires.[6] As a result, a significant increase in the level of internal and external trade took place in Buenos Aires and its area of influence.

During most of the early colonial period, Buenos Aires maintained a trade deficit with Spain. This deficit resulted in a constant outflow of gold and silver (in the form of coins, bars, and silverware).[7] Coins came to Buenos Aires from the regions of Upper Peru (the area occupied today by Bolivia and Peru). The Royal Villa of Potosí was the major center of economic activity in Upper Peru, with its adjacent silver mines and the regional mint house. Smuggling of European linen and relatively inexpensive Brazilian products was common in the port of Buenos Aires. Most of these products were sent to Upper Peru to be sold in exchange for gold and silver coins. The proceeds, especially high-quality coins, were then exported to Europe via Buenos Aires.

The constant outflow of coins from the port of Buenos Aires was perceived as creating significant liquidity problems in the area of the Rio de la Plata (Elía 1942a, 420–21). For example, as a result of numerous local complaints, in October 1618 the King of Spain passed a resolution allowing these colonies to use "products of the land" (instead of gold and silver coins) to pay the "Indies taxes" (Elía 1942a, 418).[8] Interestingly, there is very similar evidence of

[6] Before 1776, the colonies of the Rio de la Plata were under the control of the representative of the king residing in the Peruvian area.

[7] Halperín Donghi (1972, 48) reports that exports of gold and silver from Buenos Aires amounted to about 80 percent of total exports in 1796.

[8] Also, in 1622, in an attempt to stop the outflow of precious metals from Upper Peru, the Spanish Crown created the customs of Córdoba, an inland post that was supposed to control all trade between the port of Buenos Aires and the highland regions in the northwest (Upper Peru). The effectiveness of this measure was undermined by rampant smuggling.

perceived shortages of specie in colonial Canada at the end of the 18th century. Redish (1984) argues that these common complaints about a *general* scarcity of specie in Canada should really be interpreted as reflecting the discomfort of merchants with what was actually a scarcity of *high-quality* coins. Redish provides evidence indicating that most of the coins in circulation in late 1700s Canada were old coins whose weight had been reduced by intentional clipping or sweating (and the normal wear and tear of very old coins), i.e., low-quality coins. The idea behind Redish's interpretation is that, in accordance with modern versions of Gresham's Law, low-quality coins tended to drive high-quality coins out of circulation.[9]

In principle, it seems that Redish's hypothesis could be applicable for interpreting the complaints about general scarcity of specie during the early colonial period of Argentina. However, more research is needed in this area before reaching a more definite conclusion. Here, though, the focus will be on trying to identify situations where the shortage could be associated exclusively to low-denomination media of exchange, and I will restrict my study to only the latter part of the colonial period (that is, since the creation of the Viceroyalty of the Rio de la Plata in 1776).

The Monetary System

The foundations of the monetary system in the Viceroyalty of the Rio de la Plata resembled those of the system in place in Spain at the time. Basically, there were in circulation gold and silver coins minted in Spain, Mexico, and Peru. The two main mints in the Spanish colonies of South America were located in Lima and Potosí. Most of the coins circulating in the territories of the Rio de la Plata (today Argentina) were silver coins minted in Potosí (Elía 1942a, 429).

The Potosí mint was under the direct control of the Spanish Crown, which held a monopoly for issuing coins. Mining, on the other hand, was a private enterprise. The Crown, though, provided miners with most of the essential inputs and heavily taxed their production. One of the most notorious institutions of the time was the *mita*, an annual recruitment of forced Indian labor that was assigned to the different miners according to a system of concessions. In 1779 the Spanish Crown created the *Banco de San Carlos* that provided credit and other basic inputs to miners in Potosí and monopolized the purchase

[9] Sargent and Velde (2002, 125) discuss the "bullion famine" of medieval Europe in the context of their model and also conclude that talking about general shortages of coins is difficult to rationalize. Instead, they maintain that the monetary anomalies involved were the consequences of shortages of small change.

Figure 1 Silver Pesos Minted in Potosí Between 1767–1770

Source: http://www.historiadelpais.com.ar/

of silver in the region (Tandeter 1992, 198–99). The *diezmo*, a 10 percent tax, was charged on all silver production.

The mint issued coins according to general orders from the Spanish Crown.[10] The relative supply of different denominations of coins was in principle determined by directives from Spain and decisions by the mint's administrators (part of the Crown's bureaucracy). Under this scheme the supply of fractional money, in principle, did not automatically adjust to its demand, and imbalances became common.[11] During the second half of the 18th century, a system of two-year concessions was instituted for the administration of the mints on behalf of the Crown. The concession contracts stipulated targets for the cost of production and the proportion of fractional money to be minted (Dargent Chamot 2005, Ch. 15).

There were two types of silver coins in circulation at the time of the creation of the Viceroyalty: the coins called *de cordoncillo* and the older, hammered coins (or cobs) called *macuquina*. The coin de cordoncillo had the edges marked to prevent clipping and the macuquina had variations in thickness, weight, and shape, making it a coin of mediocre quality. The macuquina had been in circulation since the time when hammering was the common minting practice (in Potosí, from 1575 until 1773). In principle, both coins de cordoncillo and macuquina of all existing denominations were in circulation.

[10] Romano (1998, 133) reports that the original 16th century royal ordinances creating the mints in Mexico and Peru, including the one in Potosí, stated explicitly the rules for the proportions of the different denomination of coins to be minted. In particular, only a fourth of the coins were supposed to be of low denomination. These original orders, though, were not always strictly followed and changed over time.

[11] In an alternative scheme called "free minting," private agents holding silver or gold are able to go to the mint and exchange their metal for coins of the denomination of their choice. This scheme was predominant in Europe (see Sargent and Velde 2002, 20).

Table 1 Silver Coins—Main Denominations

Cuartillo	1/4 real
Medio Real	1/2 real
Real	
Real de a Dos	2 reals
Medio Peso	4 reals
Peso	8 reals

However, a high proportion of the stock of circulating low-denomination coins were macuquina, as these were the coins that had been minted for a longer period of time.[12]

The high-denomination, full-bodied silver coins were commonly called *plata doble* and the low-denomination coins, *plata sencilla* (Bonura 1992, 40; Tandeter 1992, 157). Most plata sencilla was of the macuquina type and not full-bodied (due to intentional clipping and the normal wear and tear resulting from their use and age). The plata doble, on the other hand, were mostly high-quality, full-bodied silver coins that were relatively scarce and especially useful for payments of imported goods from Spain.

In terms of the denomination structure, the main denominations were the *peso*, also called *peso fuerte* and the *real* with a nominal value of one-eighth of the peso. There were also coins of half, two, and four reals (Elía 1942a, 432). *Cuartillos,* coins of one-fourth of a real, were only minted in Potosí after 1794 and in very small quantities (Dargent Chamot 2005, Ch. 17).

Gold coins were very scarce in the area and of relatively high purchasing power. The most common gold coin was the doubloon of eight, which was equivalent to approximately 16 silver pesos and was mostly used for international trade and hoarding. Overall, gold coins were not used in small domestic transactions, and for this reason they play no major role in the discussion that follows.

In 18th century Spain, small change was partly provided by the issuing of *vellón*, a low-denomination coin made with a mixture of copper and small quantities of silver. According to Bonura (1992, 39–40), the vellón did not circulate in the region of the Rio de La Plata (see also, Cortés Conde and McCandless 2001, 384). These token coins were not commonly minted in Potosí, probably because of their high minting costs and an alleged (yet, somewhat surprising) reluctance of the general population to accept them in exchange. Interestingly, Romano (1998, 133) reports extensively on a similar

[12] Low-denomination coins were minted in relatively low proportions and, hence, most of the stock in circulation was fairly old. High-denomination coins were minted more intensively and were exported in large proportions, resulting in a stock with a lower average age.

phenomenon taking place in colonial Mexico (see also Hamilton 1944, 35). The reason for this phenomenon, however, is not yet well understood.[13]

The Problem of Small Change

On several occasions between 1770 and 1810, the local elites complained to the Crown about the shortage of small change. In 1773 the king, partly as an attempt to deal with the scarcity of small change, banned any export of fractional money from the colonies. Specifically, the king prohibited the shipment of pieces of half, one, and two reals to Spain and instructed the Viceroys to intensify their efforts to ensure that the royal mints coin enough silver in those denominations "for the vast commerce of America" (Hamilton 1944, 37).[14]

The macuquina was heavily used in domestic transactions, usually circulating at par value. However, its poor quality complicated its normal use and generated many complaints among the locals.[15] Its irregular shape made the macuquina very susceptible to clipping, creating uncertainty about its intrinsic value. Furthermore, by 1784 all circulating macuquina was at least ten years old (hammered coins were last produced in Potosí around 1773) and, hence, in very bad shape. At that time, the king issued an order to collect and remint all the macuquina in the colonies. After five years, in 1789 the order was reissued, allowing for a fixed two-year period to complete the process. In fact, after those two years, Viceroy Arredondo again postponed the recovery period with no explicit time limit. This process suggests that the officials in the colonies were reluctant to enforce the order to remint the macuquina as they perceived that doing so would only aggravate the shortage of small change. In fact, during the same period, Viceroy Arredondo proposed the creation of a token coin to be used in domestic trade. He explicitly pointed to the scarcity of small change as a justification. The Spanish Crown denied the proposal and instead ordered that all mint houses in the area start minting cuartillos (Elía 1942a, 425–26; Dargent Chamot 2005, Ch. 15).

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[13] Vellón was issued in colonial Mexico in 1542 and the general population refused to accept it. The experiment was a complete failure. Hamilton (1944, 36) argues that the vellón was so grossly overvalued that the population preferred to continue using cocoa beans as a mean of payment. Romano (1998, 134–35), instead, suggests that the local elites actually opposed the issuing of vellón based on political motivations. The reason vellón was not issued in the Spanish colonies in America is an unanswered question in the literature.

[14] Butlin (1953, 81–82) reports that the government in Australia (a British colony) used similar legal instruments to avoid the export of coins in 1813. Sargent and Velde (2002) report legal restrictions on the export of coins in medieval England (p. 132) and in Venice on May 1268 (p. 163). See Wallace and Zhou (1997) for a rationalization of this type of official restriction.

[15] See Hamilton (1944, 25–26) for an account of similar problems in the area of colonial Mexico.

The nature of the small change problem was twofold. First, low-denomination coins (that is, coins of two reals or less) were minted in significantly smaller proportions than the silver pesos; and second, the purchasing power of the lowest-denomination coin was very high.[16]

In terms of the relative amounts of low-denomination coins that were minted in Potosí, Tandeter (1992, 157) reports that, in the mid-1700s, 85 percent of the minting was done in plata doble (i.e., coins of four reals and higher). By the same token, Romano (1998, 117) explains that most of the coins minted in the Spanish colonies in America during the 18th century were of high denomination, with the eight-real (peso) pieces amounting to at least 95 percent of the annual issues of silver coins both in Mexico and in Peru.[17] Dargent Chamot (2005, Ch. 15) reports the reluctance of the Superintendent of the mint of Potosí in 1784 to issue large quantities of cuartillos, considering them too costly to produce. In general, cuartillos were issued in very limited amounts, and only later in the period (in Potosí, starting only in 1794).[18]

One way to get an idea of the high purchasing power of the denomination structure that was predominant in the region is to compare the value of silver coins with the level of nominal wages for unskilled rural workers. For example, at the time, a slave in the rural areas near Buenos Aires would normally receive an allowance of one real per week to buy "soap and tobacco." A free rural seasonal worker (a peon) had an average wage of around four pesos per month (although monthly wages fluctuated significantly across workers, from two to seven pesos; see Amaral 1987, 267–72). This monthly wage implied a daily wage of around one-and-a-half reals that amounted to three coins of half real, which was effectively the smallest denomination coin. A similar situation took place in the early stages of other monetary systems. For example, Hanson (1979, 283) reports that, of the common coins in circulation in Pennsylvania and Massachusetts (both British colonies) in 1742, the lowest-denomination coin represented about three days' wages for an unskilled laborer at the time.

Some Consequences

The lack of small-denomination coins resulted in the use of unofficial means of payment in everyday transactions (Bonura 1992, 40). One of these instruments, the *contraseñas*, became very popular. Contraseñas were small metal (tin) discs with the initials of the issuer printed on them (Elía 1942a, 428;

[16] Bonura (1992, 39) recognizes the relatively high purchasing power of the cuartillo and finds it puzzling that no lower-denomination coins were issued before 1794.

[17] The proportion of low-denomination coins *in circulation* was probably higher since high-denomination coins were more intensively exported.

[18] Starting in 1793, cuartillos were also minted in colonial Mexico. According to Hamilton (1944, 38), the cuartillos were "too small for convenient use and struck in inadequate quantities" and, hence, "did not end the disorder in the fractional coinage."

Prebisch 1921, 199). In everyday transactions requiring small change that the parties (buyer and seller) lack, the buyer could make payments in two possible ways. One way was to pay using contraseñas previously issued by the particular merchant participating in the transaction, in which case the transaction would terminate with the payment. The other way was for the buyer to pay in high-denomination silver coins. In this case, when necessary, the change resulting from the transaction would be provided in contraseñas issued by the seller. Sometimes, even the contraseñas issued by a third party were used as change. In general, the third party was a well-known merchant in the area and the individuals engaged in the transaction were holding his contraseñas as a result of previous transactions.[19]

The use of contraseñas did not, of course, solve all the problems. In fact, their extensive use resulted in widespread fraud and falsification. Later on, contraseñas were gradually substituted with private IOUs issued directly on paper.[20] These IOUs were inconvertible and also circulated widely in the region (Prebisch 1921, 199). They are a precursor of the inconvertible paper money that was introduced in the region more than a decade after the 1810 Revolution.[21]

Another way people circumvented the lack of small change was by developing even simpler credit arrangements. Customers would build up a debit at the community store until it was possible to settle the payment using higher-denomination coins that were more readily available (Schmit 2003, 265). Obviously, the use of this kind of informal credit was limited to cases where the owner of the store was relatively certain that the customer had reasons to secure a permanent relationship with the store.

The Premium

There is some evidence that in the City of Buenos Aires during colonial times, the hard peso sometimes circulated at a premium over fiduciary silver coins, i.e., the low-denomination, usually not full-bodied plata sencilla (Prebisch

[19] In colonial Mexico, it was popular to use for payments small wooden disks with the name of the issuer (a merchant) printed on them. These disks were called *tlacos* and they emerged in response to the recurrent shortages of small change that took place in the Mexican territory during the 18th century (Romano 1998, 137; Hamilton 1944, 36–38). Tin-made tokens with similar characteristics as the contraseñas circulated in England in 1576 (Sargent and Velde 2002, 266).

[20] Butlin (1953, 26–27) describes privately issued *promissory notes* that circulated in Australia during the colonial period and the rampant forgery that originated around them.

[21] Hanson (1979, 285) convincingly argues that the origin of paper currency in the colonies of North America was the result of the persistent shortages of low-denomination coins. He provides evidence of the issuance of private circulating notes by merchants early in the process. Sargent and Velde (2002, 203) discuss evidence from 1577 France that documents the widespread use of private IOUs in response to the shortages of small change.

1921, 195; Bonura 1992, 40–41).[22] Interestingly, the premium was lower (and even zero) in the periphery (the *interior*), creating a flow of plata sencilla from Buenos Aires to those regions.[23] In 1790 the authorities in Buenos Aires asked the Crown to introduce legal restrictions to abolish "the 3 percent premium of the hard peso." In 1798, after the Crown did not respond, the request was reiterated. The main justification for the request was the constant flow of fractional coins out of Buenos Aires to the interior. The official document stated that this flow had "reduced the quantity of small-denomination coins, ... creating difficulties in the change or reduction of the plata doble to the sencilla ... the specie so necessary for making small daily purchases, which are very indispensable transactions" (see Bonura 1992, 41).[24]

With respect to the evolution of the premium over time, it appears that the premium was fairly constant. Bonura (1992, 49), for example, reports that the premium was still around 3 percent in 1812, when the authorities in Buenos Aires engaged in another legal attempt to reduce it. Sargent and Velde (2002) associate periods of shortages of small change with periods of depreciation in the value of fractional money. The evidence from Argentina is too sparse to test this hypothesis (but, in principle, no clear trend in the premium was observed in the region).

2. EVIDENCE FROM EARLY ARGENTINA

In 1810, the *Cabildo* (the town council) of Buenos Aires declared autonomy from the Spanish Crown. With the end of the Viceroyalty of the Rio de la Plata, Buenos Aires lost Upper Peru from its area of influence; and with Upper Peru, the mint of Potosí and the silver mines.[25]

This transitional period was associated with general monetary disarray in the region. The confrontation with Upper Peru (which had remained loyal to the Crown) and the necessary financing of military expenses (including significant imports that needed to be paid in specie) created a sharp contraction in the amount of available means of payment in Buenos Aires (Prebisch 1921, 198). During this period, many government officials proposed a compulsory

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[22] Tandeter (1992, 157) reports that the plata doble had a premium over the plata sencilla in the Villa of Potosí in the mid-1700s. He attributes the premium to the fact that the plata doble was the one preferred in long-distance trade.

[23] This kind of geographic dispersion in the exchange rate of coins was also observed across the French territory during the 1570s, a period of monetary "chaos" (Sargent and Velde 2002, 200).

[24] In principle, one would expect shortages of small change to be associated with, if anything, a premium on low-denomination coins. This is the opposite of what is reported here. It seems likely, however, that the premium was not uniform across transactions, and that the 3 percent premium on hard pesos was predominant only in large-value transactions and international trade.

[25] For a good overview of the economic factors that led to the breakup of the Viceroyalty into different countries, see Cortés Conde and McCandless (2001).

remint of all the old coins in circulation. Furthermore, some of these proposals included the imposition of a steep proportional tax upon reminting. In view of these proposals, it seems that hiding and hoarding coins was a natural reaction of the population.

The new government made several attempts at issuing new coins during this period. In 1813, after temporarily recovering the city of Potosí, the first Argentinean coin was minted. A year later, however, the Independence Army lost Potosí to royal forces and the minting stopped.[26] Some minting of silver pesos took place in the province of Córdoba during 1815, but in very limited amounts (Elía 1942a, 433). At that time, illegal private minting of cobs and other (very low quality) silver counterfeits was common in the northwest region of the country (Bonura 1992, 73). In 1817 Governor Güemes officially authorized the circulation of "illegal" coins (after being officially stamped) in the territory under his jurisdiction in the northwest part of the country. He gave as a justification for this resolution the "evils associated with the lack of means of payment" (Elía 1942a, 435). A year later, the federal authority banned the circulation of these "Güemes" coins, establishing a severe punishment for those who accepted and/or held them. Overall, no real progress was made in providing the economy with appropriate means of payment during the first decade after the revolution (Bonura 1992, 81).

Besides the general monetary disorder, some specific episodes suggest the existence of shortages of fractional money. In this respect, two situations appear most relevant: the provision of copper coins approved in 1821 after several years of discussions and the authorization granted to the Bank of Buenos Aires to issue paper notes of relatively low denomination in 1823.[27]

Copper Coins

In June 1815 the newly created government in Buenos Aires started evaluating the introduction of "provisional money" in the form of copper coins (Bonura 1992, 61). For this purpose, the government commissioned an extraordinary consulting body of experts to study the issue. The authority's motivation for the introduction of these coins was twofold. The first was that shortages of small change were a recurrent problem that needed to be fixed. In August 1815 the body of experts presented a detailed report in which they unanimously agreed

[26] Minting of Argentinean coins in Potosí resumed for a short period in 1815.

[27] Experiences in other countries influenced these two decisions. For example, the general perception in Buenos Aires was that copper coins were being used with great success in Portugal (Bonura 1992, 77). With respect to banking, reports of the benefits associated with the operation of the Bank of England were one of the main motivations for the creation of the Bank of Buenos Aires (Prebisch 1921, 199–200).

Figure 2 Copper Coin Minted at Boulton's Mint

http://www.camoar.gov.ar/CecasProvinciales.htm

that introducing copper coins was essential for eliminating the inconveniences resulting from the persistent lack of small change (Bonura 1992, 65).

The second motivation was the possibility of obtaining extra resources for a government that was in desperate need of financing.[28] This issue was the subject of important disagreement among experts. They discussed the estimated costs of minting copper coins extensively, but they did not reach an agreement, so the implementation was postponed (Bonura 1992, 65). Sporadically, during the next five years, the government authorities in Buenos Aires revisited the possibility of issuing copper coins but never managed to implement the idea.[29]

Finally, in October 1821, a law was passed allowing the government to arrange the minting of 100,000 pesos in copper coins of one-tenth of a real (Elía 1942a, 437). These coins, the first Argentinean copper coins, were minted in Birmingham, England (at Boulton's mint). Fifty thousand pesos of those coins were received and put into circulation in July 1823.[30] Elía (1942a,

[28] Token coins are usually circulated at a value greater than their intrinsic value. For this reason, they have the potential to become a source of revenue for the monetary authority. During 1817, the government was evaluating the possibility of opening a mint in the city of Córdoba. After concluding that the project would not be profitable for the government, the idea was abandoned. The evidence seems to indicate that there was a problem of insufficient scale of production. Apparently, the set-up costs of operating a mint were very high, and the quantity of metal available from the mines of Famatina, the planned source of basic input, was not enough to make the enterprise profitable (Bonura 1992, 72).

[29] The issue of minting copper coins was again extensively discussed in 1818 when the new government was evaluating the possibility of establishing an official mint in Buenos Aires (Bonura 1992, 75).

[30] To put some perspective on these numbers, note, for example, that in 1923 total tax revenue for the province of Buenos Aires was around two million pesos (see Bordo and Vegh 2002,

437) reports that the public immediately absorbed that first lot of coins and the government then requested that the Birmingham mint deliver the rest of the coins as soon as possible. Both the small denomination of these copper coins and their generalized acceptance by the public seem indicative of the high level of unsatisfied demand for fractional money that existed during that period.[31]

Paper Money

In June 1822 the government in Buenos Aires gave a group of local businessmen an exclusive 20-year concession to create the first (and only) bank in the region. The Bank of Buenos Aires (also called *Banco de Descuentos*) was supposed to be fully funded with private capital. Part of the government's justification for allowing the creation of the Bank was the need to provide appropriate means of payment to the community (Irigoin 2003, 65).[32] Trade liberalization after the revolution resulted in a substantial increase in commercial activity, in turn creating the urgent need for more developed monetary and financial institutions in the region.

However, this was not the only motivation for the creation of the Bank. In fact, there is some evidence indicating that the primary reason was to allow the government to access cheaper financing. By 1822 the government was heavily involved in a civil war and was quickly running out of resources (Prebisch 1921, 201). The plan was that the government would take loans from the bank at preferential rates.

Some of the factors that triggered the creation of the Bank of Buenos Aires seem indicative of the persistent shortage of low-denomination money. First, from discussions at the time it is clear that private IOUs (*vales*) and contraseñas were still in circulation when the Bank was created in 1822 (Elía 1942b, 323; Prebisch 1921, 199; Irigoin 2003, 65). The use of vales and contraseñas can be taken as evidence of the need for fractional money. To the extent that parties in a transaction were willing to accept these very imperfect means of payment, which were clearly associated with significant risk of fraud

Table 1). In other words, the gross revenue from the introduction of this first batch of copper coins in 1823 was about 2.5 percent of total annual tax revenue.

[31] The circulation of token coins was never automatic in the early stages of monetary development. It often happened that the population, accustomed to full-bodied coinage, distrusted the validity of token copper coins as an acceptable means of payment. See Butlin (1953, 37) for the case of Australia and Sargent and Velde (2002, 210) for the case of France in the 1590s (see also footnote 13 in this paper). Of course, counterfeiting was always a potential problem in the case of circulating token coins (see, for example, Sargent and Velde 2002, 217–18). The fact that these first copper coins were minted in England using frontier technology at the time probably reduced the risk of counterfeiting making the coins more likely to circulate.

[32] Redish (1984) reports that a similar justification was used at the time of the creation of the first Canadian banks at the beginning of the 19th century.

and counterfeiting, it must be the case that no better payment methods were available (Prebisch 1921, 199).

Second, while initially the Bank of Buenos Aires was only allowed to issue notes in denominations no lower than 20 hard pesos, in mid-1823 the government agreed to authorize the Bank to start issuing lower-denomination bills. These bills came to replace some Treasury notes of similar denomination that the Ministry of Finance had introduced only months before. The Bank issued bills of one, three, and five hard pesos, convertible to gold and silver coins upon presentation at the Bank's window (Elía 1942b, 326). While these denominations were not, by any means, the lowest of the prevailing structure, they were commonly used in domestic transactions. Also, they were probably considered the natural intermediate step in the move toward lower denominations.

During the first two years of its existence, the Bank issued convertible money notes well in excess of its reserves of gold and silver that resulted in a confidence crisis in 1825. Early in 1826, the Bank was taken over by the government and its money notes were declared inconvertible. The notes stayed in circulation but only based upon government fiat. The perceived insufficiency of means of payment prevailing in the region was then replaced by excessive printing of inconvertible paper money. A regime of high inflation followed, which lasted for many decades.[33]

3. FINAL REMARKS

In this article I reviewed evidence that suggests that shortages of small change were a problem in the economy of the Rio de la Plata area during the colonial period and the first two decades after independence. Evidence of this sort is already available for several other regions around the world. It is interesting

[33] After 1825 the authorities in both Bolivia and Buenos Aires started a period of sustained monetary expansion and inflation (Irigoin 2003, 60). In Bolivia, the minting and systematic debasement of silver coins (*moneda feble*) was the main fiscal instrument of the new government. These coins circulated also in the northwest regions of Argentina. In Buenos Aires, the government printed large amounts of inconvertible paper money to finance increasing fiscal deficits. The paper peso depreciated around 200 percent during 1826 and continued depreciating in the following years. For a detailed discussion of the monetary history of Argentina during this period, see Irigoin (2003), Bordo and Végh (2002), and Irigoin (2000). In general, the paper money from Buenos Aires did not circulate in the provinces. In the interior of the country, several provincial governments attempted to issue their own paper money but faced substantial problems in inducing its circulation, as the general population deeply mistrusted the viability of fiat money. In 1826 the province of Corrientes issued 3,000 hard pesos in low-denomination notes, but acceptance was limited and the experiment became a complete failure (Irigoin 2003, 67). Sometimes the government introduced extreme legislation to try to encourage the circulation of the money notes. For example, in 1840 the provincial authority in Tucumán instituted the death penalty for those not accepting in exchange the paper money printed by the Northern League, a coalition of northwestern provinces (Halperín Donghi 1979, 91).

to verify that a similar monetary phenomenon took place in Argentina during the early stages of its political and economic development.

Several features made the evidence presented here especially interesting. First, while most of the European evidence comes from economies where free minting was in place, in Argentina during the late colonial period the supply of coins was under the direct control of the Spanish Crown. Minting policies, then, were not uniquely directed to improve the smooth functioning of the monetary economy in the colonies. Spain was the main provider of high-quality silver coins to the rest of Europe during that time. For this reason, a major motivation for the Crown's policies was to maintain an international reputation of high quality for Spanish coins. These competing objectives probably increased the chances of misalignments between demand and supply of denominations.

Second, the evidence clearly illustrates the interaction between money and credit during the early stages of monetary development. To bypass the problem of small change, agents in the economy developed rudimentary credit arrangements that allowed them to trade with one another. Two schemes were prevalent. In one scheme, the buyer would extend credit to the seller through the use of contraseñas; in the other, the seller would grant credit to the buyers by allowing them to accumulate a debit in a temporary account. It is a general principle in monetary economics that money and credit act as close substitutes. In general, however, the emphasis has been on explaining how monetary exchange increases the trading possibilities in an economy where credit is not always feasible (as in Kiyotaki and Wright 1993). The evidence presented here highlights the reciprocal fact that when the convenience of monetary exchange is undermined by, in this case, the lack of small change, agents turn to imperfect credit arrangements to carry out their economic transactions. (See Jin and Temzelides [2004] and Cuadras-Morató [2005] for a formal discussion of some of these issues.)

Finally, it was interesting to see the newly created government confronting all the basic economic issues involved in the provision of small change when deciding to introduce copper coins. On one hand, effective fractional coins needed to be of relatively high quality to avoid counterfeiting. On the other hand, high-quality, low-denomination coins were very costly to produce. The government realized that only a large scale of production could lower the unitary cost of production to an acceptable level. The lack of sufficient mineral input delayed production of copper coins for several years. In the end, the government resorted to importing the coins from England, an international producer of coins for which scale of production was obviously not an issue.

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Implementation of Optimal Monetary Policy

Michael Dotsey and Andreas Hornstein

Recently the study of optimal monetary policy has shifted from an analysis of the welfare effects of simple parametric policy rules to the solution of optimal planning problems. Both approaches evaluate the welfare effects of monetary policy in an explicit monetary model of the economy, but they differ in the scope of analysis. The first approach is more restrictive in that it finds the optimal policy within a class of prespecified policy rules for the monetary policy instrument. On the other hand, the second approach finds the optimal monetary policy among all allocations that are consistent with a competitive equilibrium in the monetary economy. Since monetary policy, in general, does not choose the economy's allocation but implements policy through a rule for the policy instruments, it is natural to ask whether the policy rule implied by the solution to the planning problem implements the optimal planning allocation. In most work on optimal planning problems, it is indeed taken for granted that the solution of the planning problem can be implemented through some policy rule for the monetary policy instrument but, as we show in this article, this need not always be the case.

There is a vast literature on optimal monetary policy that studies the solution to planning problems. The environments examined are diverse, ranging from models in which there are no private sector distortions other than an inflation tax to models where economies are subject to various types of nominal rigidities. The policymaker is assumed to choose among all the allocations that are consistent with a market equilibrium in the given environment. In addition, different assumptions are made as to whether a policymaker can or cannot commit to his future choices. Under a full-commitment policy, we as-

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sume that the policymaker chooses all current and future actions in an initial period. Alternatively, under time consistency we assume that in every period a policymaker chooses the optimal action, taking past outcomes as given. For either specification, the solution to the planning problem specifies a rule that determines the allocation, and part of the allocation is the setting of the policy instrument.

The question is whether the policy rule implied by the solution to the planning problem (or a variation thereof) can implement the optimal allocation for the planning problem. Specifically, how would the competitive economy behave if the monetary authority simply announced the policy rule implied by the solution to the planning problem? In particular, conditional on the policy rule, will there be a unique competitive equilibrium?

Giannoni and Woodford (2002a, 2002b) discuss the implementability of optimal policy for local approximations of the planning problem with full commitment. This starts with a log-linear approximation around the steady state of the solution to the full-commitment problem. Within the approximation framework, implementability of the optimal policy rule is equivalent to the existence and uniqueness of rational expectations equilibria in linear models. As such, implementability is concerned with "dynamic" uniqueness, that is, the existence of a unique stochastic process that characterizes the competitive equilibrium.

King and Wolman (2004) discuss the implementation of Markov-perfect policy rules for time-consistent solutions to the planning problem. King and Wolman (2004) show that Markov-perfect policies with an optimal nominal money stock instrument can imply equilibrium indeterminacy at two levels. First, it can imply multiple steady states. Second, around each steady state it can imply static price level indeterminacy, that is, conditional on future outcomes there can be multiple current equilibrium prices.

In this article, we review implementability of both the optimal full-commitment and time-consistent Markov-perfect monetary policies when the policymaker uses a nominal money stock instrument. We study optimal policy in a simple New Keynesian economic model as described in Wolman (2001) and King and Wolman (2004). We first characterize the solution to a linearized version of the first-order conditions (FOCs) of the planning problems. We show that optimal monetary policy locally implements the planning allocation for the full-commitment and the Markov-perfect case. We then study whether the policy rules implement the planning allocations globally. We review King and Wolman's (2004) argument that the Markov-perfect policy rule cannot implement the planning allocation. Finally, we provide a partial argument that the full-commitment policy rule globally implements the planning allocation.

1. A SIMPLE ECONOMY WITH STICKY PRICES

We investigate the question of the implementability of optimal monetary policy within the confines of a simple New Keynesian economic model. The model contains an infinitely lived representative household with preferences over consumption and leisure. The consumption good is produced using a constant-returns-to-scale technology with a continuum of differentiated intermediate goods. Each intermediate good is produced by a monopolistically competitive firm with labor as the only input. Intermediate goods firms set the nominal price for their products for two periods, and an equal share of intermediate firms adjust their nominal price in any period. We describe a symmetric equilibrium for the economy, and we characterize the two distortions that make the equilibrium allocation suboptimal relative to the Pareto-optimal allocation.

The Representative Household

The representative household's utility is a function of consumption, c_t, and the fraction of time spent working, n_t,

$$E_0 \sum_{t=0}^{\infty} \beta^t \left[\ln c_t - \chi n_t\right], \tag{1}$$

where $\chi \geq 0$, and $0 < \beta < 1$. The household's period budget constraint is

$$P_t c_t + B_{t+1} + M_t \leq W_t n_t + R_{t-1} B_t + M_{t-1} + D_t + T_t, \tag{2}$$

where P_t (W_t) is the money price of consumption (labor), B_{t+1} (M_{t+1}) are the end-of-period holdings of nominal bonds (money), R_{t-1} is the gross nominal interest rate on bonds, T_t are lump-sum transfers, and D_t is profit income from firms owned by the representative household. The household is assumed to hold money in order to pay for consumption purchases

$$M_t = P_t c_t. \tag{3}$$

We will use the term "real" to denote nominal variables deflated by the price of consumption goods, and we use lower-case letters to denote real variables. For example, real balances are $m \equiv M/P$.

The FOCs of the representative household's problem are

$$\chi = w_t / c_t, \text{ and} \tag{4}$$

$$1 = \beta E_t \left[\frac{c_t}{c_{t+1}} \cdot \frac{R_t}{P_{t+1}/P_t} \right]. \tag{5}$$

Equation (4) states that the marginal utility derived from the real wage equals the marginal disutility from work. Equation (5) is the Euler equation, which states that if the real rate of return increases, then the household increases future consumption relative to today's consumption.

Firms

The consumption good is produced using a continuum of differentiated intermediate goods as inputs to a constant-returns-to-scale technology. Producers of the consumption good behave competitively in their markets. There is a measure one of intermediate goods, indexed $j \in [0, 1]$. Production of the consumption good c as a function of intermediate goods, $y(j)$, used is

$$c_t = \left[\int_0^1 y_t(j)^{(\varepsilon-1)/\varepsilon} dj \right]^{\varepsilon/(\varepsilon-1)}, \tag{6}$$

where $\varepsilon > 1$. Given nominal prices, $P(j)$, for the intermediate goods, the nominal unit cost and price of the consumption good is

$$P_t = \left[\int_0^1 P_t(j)^{1-\varepsilon} dj \right]^{1/(1-\varepsilon)}. \tag{7}$$

For a given level of production, the cost-minimizing demand for intermediate good j depends on the good's relative price, $p(j) \equiv P(j)/P$,

$$y_t(j) = p_t(j)^{-\varepsilon} c_t. \tag{8}$$

Each intermediate good is produced by a single firm, and j indexes both the firm and good. Firm j produces $y(j)$ units of its good using a constant-returns technology with labor as the only input,

$$y_t(j) = \xi_t n_t(j), \tag{9}$$

and ξ_t is a positive iid productivity shock with mean one. Each firm behaves competitively in the labor market and takes wages as given. Real marginal cost in terms of consumption goods is

$$\psi_t = w_t/\xi_t. \tag{10}$$

Since each intermediate good is unique, intermediate goods producers have some monopoly power, and they face downward sloping demand curves, (8). Intermediate goods producers set their nominal price for two periods, and they maximize the discounted expected present value of current and future profits:

$$\max_{P_t(j)} \left(\frac{P_t(j)}{P_t} - \psi_t \right) y_t(j) + \beta E_t \left[\frac{c_t}{c_{t+1}} \cdot \left(\frac{P_t(j)}{P_{t+1}} - \psi_{t+1} \right) y_{t+1}(j) \right]. \tag{11}$$

Since the firm is owned by the representative household, the household's intertemporal marginal rate of substitution is used to discount future profits. Using the definition of the firm's demand function, (8), the first-order condition for profit maximization can be written as

$$
\begin{aligned}
0 &= \left(\frac{P_t(j)}{P_t}\right)^{1-\varepsilon}\left(1-\mu\frac{\psi_t}{P_t(j)/P_t}\right) \\
&\quad +\beta E_t\left[\left(\frac{P_t(j)}{P_{t+1}}\right)^{1-\varepsilon}\left(1-\mu\frac{\psi_{t+1}}{P_t(j)/P_{t+1}}\right)\right],
\end{aligned}
\tag{12}
$$

with $\mu = \varepsilon/(\varepsilon - 1)$.

A Symmetric Equilibrium

We will assume a symmetric equilibrium, that is, all firms who face the same constraints behave the same. Each period, half of all firms have the option to adjust their nominal price. This means that in every period there will be two firm types: the firms who adjust their nominal price in the current period, type 0 firms with relative price p_0, and the firms who adjusted their price in the last period, type 1 firms with current relative price p_1.

Conditional on a description of monetary policy, the equilibrium of the economy is completely described by the sequence of marginal cost, relative prices, inflation rates, nominal interest rates, aggregate output, and real balances $\{\psi_t, p_{0,t}, p_{1,t}, \pi_t, R_t, c_t, m_t\}$ such that (3), and

$$\psi_t = \chi c_t/\xi_t, \tag{13}$$

$$1 = \frac{1}{2}\left[p_{0,t}^{1-\varepsilon} + p_{1,t}^{1-\varepsilon}\right], \tag{14}$$

$$0 = p_{0,t}^{1-\varepsilon}\left(1-\mu\frac{\psi_t}{p_{0,t}}\right) + \beta E_t\left[p_{1,t+1}^{1-\varepsilon}\left(1-\mu\frac{\psi_{t+1}}{p_{1,t+1}}\right)\right], \tag{15}$$

$$\pi_{t+1} = \frac{p_{0,t}}{p_{1,t+1}}, \text{ and} \tag{16}$$

$$1 = \beta E_t\left[\frac{c_t}{c_{t+1}}\cdot\frac{R_t}{\pi_{t+1}}\right]. \tag{17}$$

Equation (13) uses the optimal labor supply condition (4) in the definition of marginal cost (10). Equation (14) is the price index equation (7) and equation (15) is the profit maximization condition (12) for the two firm types. Equation (16) just restates how next period's preset relative price $p_{1,t+1}$ is related to the relative price that is set in the current period, $p_{0,t}$, through the inflation rate π_{t+1}. Finally, equation (17) is the household's Euler equation, (5).

Distortions

Allocations in this economy are not Pareto-optimal because of two distortions. The first distortion results from the monopolistically competitive structure of intermediate goods productions: the price of an intermediate good is not equal to its marginal cost. The average markup in the economy is the inverse of the real wage, P_t/W_t, that is, according to equation (10), the inverse marginal cost, $1/\left(\xi_t\psi_t\right)$. The second distortion reflects inefficient production when relative prices are different from one. Using the expressions for the production of final goods and the demand functions for intermediate goods, (6) and (8), we can obtain the total demand for labor as a function of relative prices and aggregate output. Solving aggregate labor demand for aggregate output, we obtain an "aggregate" production function

$$d_t c_t = \xi_t n_t \text{ with } d_t \equiv (1/2)\left(p_{0,t}^{-\varepsilon} + p_{1,t}^{-\varepsilon}\right). \tag{18}$$

Given the symmetric production structure, equations (6) and (9), efficient production requires that equal quantities of each intermediate good are produced. Allocational efficiency is reflected in the term $d_t \geq 1$. The allocation is efficient if $p_{0,t} = p_{1,t} = d_t = 1$.

For the following analysis of optimal policy, it is useful to rewrite the household's period utility from the equilibrium allocation as a "reduced form" utility function of the markup and efficiency distortion. Combining expression (13) for equilibrium consumption as a function of marginal cost and productivity with the characterization of the aggregate production function (18) yields equilibrium work effort

$$n_t = d_t\psi_t/\chi. \tag{19}$$

We can substitute expressions (13) and (19) for consumption and work effort in the household's utility function and obtain the reduced form utility function

$$E_0\sum_{t=0}^{\infty}\beta^t\left[\ln\left(\psi_t\right) - d_t\psi_t\right], \tag{20}$$

after dropping any constant or additive exogenous terms.

2. MONETARY POLICY

Since the allocation of the above-described monopolistically competitive equilibrium with sticky prices is suboptimal, there is the potential for welfare-improving policy interventions. In view of the role of nominal rigidities, we want to characterize optimal monetary policy. In particular, we want to know how optimal monetary policy can be implemented given some choice of policy instrument. We examine the implications of choosing the nominal money stock as the policy instrument. This is the policy instrument considered in King and Wolman (2004), where they assume that the policymaker chooses a sequence for the nominal money stock $\{M_t\}$. Alternatively the policymaker could select the nominal interest rate, R_t, as the policy instrument. The choice of policy instrument can be crucial for questions of the implementability of optimal monetary policy, and we will get back to this issue in the conclusion.

For the analysis of the monetary policy planning problem, it is convenient to define monetary policy in terms of the money stock normalized relative to the preset nominal prices,

$$m_{1t} = \frac{M_t}{P_{1,t}}, \tag{21}$$

rather than the nominal money stock, M_t, directly. This normalization is not restrictive for the analysis of a policymaker that can commit to future policy choices, the full-commitment case. In the case of time-consistent policies, when a policymaker cannot commit to future policy choices, we will argue that for the particular class of Markov-perfect policies that we study, the normalized money stock is the relevant choice variable. Combining the policy rule with the cash-holding condition, (3), and using $P_{1,t} = P_{0,t-1}$, we obtain an equilibrium condition for consumption

$$c_t = p_{1,t} m_{1t}. \tag{22}$$

Optimal Monetary Policy

The objective of monetary policy is well-defined: the policymaker is to choose an allocation that maximizes the representative household's utility subject to the constraint that the allocation can be supported as a competitive equilibrium. For our simple example, any allocation that satisfies equations (13)–(16), (18), and (22) is a competitive equilibrium. We summarize these constraints as

$$E_t \left[h\left(x_{t+1}, x_t; \xi_{t+1}, \xi_t \right) \right] = 0 \text{ for } t \geq 0. \tag{23}$$

The vector $x_t = (y_t, z_t)$ contains the private sector variables, $y_t = \left(p_{0,t}, p_{1,t}, \pi_t, \psi_t, d_t, c_t \right)$, and the policy instrument, $z_t = m_{1t}$.[1] Formally, the policymaker's optimization problem is then defined as

$$\max E_0 \sum_{t=0}^{\infty} \beta^t u\left(x_t \right) \text{ s.t. } E_t \left[h\left(x_{t+1}, x_t; \xi_{t+1}, \xi_t \right) \right] = 0 \text{ for } t \geq 0, \tag{24}$$

where u denotes the period utility function of the representative household as defined in equation (20). A solution to this problem will have x_t as a function of the current and past state of the economy.

We will solve two alternative versions of the planning problem. First, we assume that the policymaker at time zero chooses once and for all the optimal allocation among all feasible allocations that can be supported as a competitive equilibrium. This approach delivers the constrained optimal allocation, but frequently the chosen allocation is not time consistent. The allocation is not time consistent in the sense that if a policymaker gets the option to reconsider his choices after some time, he would want to deviate from the initially chosen

[1] The characterization of the private sector involves equilibrium prices and quantities. With some abuse of standard terminology, we will call the vector y the equilibrium allocation.

path. The alternative approach then finds optimal time-consistent monetary policies. In particular, we will restrict attention to Markov-perfect policy rules, that is, rules that make policy choices contingent on payoff-relevant state variables only.

For the planning problem, we are not specific about how the policymaker can implement the policy: we simply assume that the policymaker can select any allocation subject to the constraint that the allocation is consistent with a competitive equilibrium allocation. We will say that a policy can be implemented if a unique rational expectations equilibrium exists when the policymaker sets the policy instrument, z_t, according to the state-contingent rule implied by the planning problem.

Optimal Policy with Full Commitment

Suppose that at time zero the policymaker chooses a sequence $\{x_t\}$ for the market allocation and the policy instrument that solves problem (24). We assume that the policymaker is committed to this outcome for all current and future values of the market outcome and the instrument. The FOCs for this constrained maximization problem are

$$
\begin{aligned}
0 &= Du(x_t) + \lambda_t E_t\left[D_2 h\left(x_{t+1}, x_t; \xi_{t+1}, \xi_t\right)\right] && (25)\\
&\quad +\lambda_{t-1} D_1 h\left(x_t, x_{t-1}; \xi_t, \xi_{t-1}\right) \text{ for } t > 0, \text{ and} \\
0 &= Du(x_t) + \lambda_t E_t\left[D_2 h\left(x_{t+1}, x_t; \xi_{t+1}, \xi_t\right)\right] \text{ for } t = 0. && (26)
\end{aligned}
$$

Note that the FOC for the initial time period, $t = 0$, is essentially the same as the FOCs for future time periods, $t > 0$, if we assume that the lagged Lagrange multiplier in the initial time period is zero, $\lambda_{-1} = 0$. This simply means that in the initial time period, the policymaker's choices are not constrained by past market expectations of outcomes in the initial period.

Marcet and Marimon (1998) show how to rewrite the planning problem as a recursive saddlepoint problem such that dynamic programming techniques can be applied. Following their approach, the Lagrange multiplier, λ_{t-1}, can be interpreted as a state that reflects the past commitments of the planner. Given the dynamic programming formulation, the optimal policy choice will then be a function of the state of the economy,

$$x_t = g_x^{FC}\left(\lambda_{t-1}, \xi_t\right) \text{ and } \lambda_t = g_\lambda^{FC}\left(\lambda_{t-1}, \xi_t\right). \quad (27)$$

The policymaker's optimization problem is not time consistent because of the particular status of the initial period. If a policymaker gets the opportunity to reevaluate his choices at some time $t' > 0$, then equation (25) will no longer characterize the optimal decision at t'. Rather equation (26) will apply at the time t', and, in general, the policymaker would want to deviate from his original decision. If the policymaker has no way to precommit to future policy actions, the optimal policy will therefore not be time consistent.

Markov-Perfect Optimal Policy

We study a particular class of time-consistent policies, namely Markov-perfect policies. For a Markov-perfect policy, the optimal policy rule is restricted to depend on payoff-relevant state variables only, that is, predetermined variables that constrain the attainable allocations of the economy. We can think of today's policymaker as taking his own future actions as given by a policy rule that makes his choices contingent on the future payoff-relevant state variables. Given these future choices, the policymaker's optimal choice for today will then also depend on payoff-relevant state variables only.

In our environment, predetermined nominal prices do not constrain the policymakers' choices among the allocations that are consistent with a competitive equilibrium. Even though the nominal price set by a firm that adjusted its price in the last period, $P_{1,t}$, is predetermined, the relevant variable is that firm's relative price, $p_{1,t}$, which is not predetermined. Since the predetermined nominal price is not payoff-relevant, the policymaker has to choose the nominal money stock in a way such that the predetermined nominal price cannot affect outcomes. But this just means that the policymaker cannot choose the nominal money stock, M_t, but has to choose the normalized money stock, m_{1t}.

Our environment as described by (23) then has the feature that, except for the exogenous shocks, ξ_t, there are no predetermined variables that constrain the equilibrium allocation. In other words, in any time period the values for the variables that characterize the competitive equilibrium have to be consistent with future values of the same variables, but the variables can be chosen independently of any values they took in the past.

In a Markov-perfect equilibrium, the current policymaker then assumes that future choices and outcomes are time invariant functions of ξ, $x_{t'} = g_x^{MP}\left(\xi_{t'}\right)$, for $t' > t$. For this reason, current policy choices have no effect on future outcomes, and the policymaker's choice problem simplifies to

$$\begin{aligned} x_t^*\left(\xi_t; g_x^{MP}\right) &= \arg\max_x u\left(x_t\right) \\ & \text{s.t. } 0 = E_t\left[h\left(g_x^{MP}\left(\xi_{t+1}\right), x_t; \xi_{t+1}, \xi_t\right)\right]. \end{aligned} \tag{28}$$

The FOCs for this problem coincide with the FOCs of the optimization problem with commitment for the initial period, equation (26).[2] In a time-consistent Markov-perfect equilibrium, the optimal policy choice satisfies $x_t^*\left(\xi_t; g_x^{MP}\right) = g_x^{MP}\left(\xi_t\right)$.

[2] In general, the FOCs for a Markov-perfect optimal policy are different from the initial period FOCs for an optimal policy with full commitment. If there are endogenous state variables, then even with Markov-perfect optimal policies, a policymaker can influence future policy choices by changing next period's state variables and thereby affecting the constraint set of next period's policymaker.

Implementability of Optimal Policy

If the only requirement for feasible monetary policy is the consistency with a competitive equilibrium, then there is no reason to distinguish between private sector choices, y_t, and the policy instrument, z_t. We might as well assume that the policymaker chooses both variables, x_t, subject to the consistency requirements. Now suppose that the outcome of the optimization problem is a policy rule that specifies choices for the instrument and the private sector allocation contingent on outcomes that may include the current and past states of the economy

$$z_t = g_{zt}(\cdot) \text{ and } y_t = g_{xt}(\cdot). \tag{29}$$

A somewhat narrower definition of what constitutes a feasible monetary policy not only requires that the allocations implied by g are consistent with a competitive equilibrium, but also requires that, conditional on the rule for the policy instrument, g_z, the rule for the private sector allocation, g_y, is the unique competitive equilibrium outcome. That is, g_y is the unique solution of

$$E_t\left[h\left(y_{t+1}, g_{z,t+1}(\cdot), y_t, g_{zt}(\cdot); \xi_{t+1}, \xi_t\right)\right] = 0 \text{ for } t \geq 0. \tag{30}$$

If we cannot find a unique solution, g_y, to this dynamic system, then we say that the optimal policy cannot be implemented since the associated competitive equilibrium is indeterminate.

In the case of full-commitment policy rules, we can consider an expanded version of the planner's policy rule. Suppose that the planner can respond contemporaneously to deviations of the competitive equilibrium allocation from the allocation implied by the full-commitment policy rule. Then we can define a modified rule for the policy instrument

$$\tilde{g}_z^{FC}\left(y_t, \lambda_{t-1}, \xi_t\right) = g_z^{FC}\left(\lambda_{t-1}, \xi_t\right) + H\left[y_t - g_{y,t}^{FC}\left(\lambda_{t-1}, \xi_t\right)\right],$$

where $H(0) = 0$. Since the choice of the function H is arbitrary, except for the origin normalization, it then appears that, under these circumstances, a planner can always implement the full-commitment solution. Note that a Markov-perfect policy rule cannot be augmented in this way since the contemporaneous private sector allocation is not a payoff-relevant state variable.

3. LOCAL PROPERTIES OF OPTIMAL POLICY

We now discuss the local dynamics of full-commitment and Markov-perfect optimal policy for our simple economy from Section 1. We derive necessary conditions for the optimal policy and characterize the deterministic steady state of the economy for the types of policy. We then study the properties of optimal policy for a local approximation around its steady state. Our approach follows King and Wolman (1999) and Khan, King, and Wolman (2003) in that we study the dynamics of a linear approximation to the FOCs and constraints of the optimal planning problem.[3] The two optimal policies imply different policy rules for a money stock instrument. We show that for the local approximation, both implied that policy rules implement a unique rational expectations equilibrium.

Consider a policymaker who uses the money supply as an instrument, that is, the policymaker chooses the money stock according to equation (21). We can then write the competitive equilibrium conditional on the instrument choice in terms of the variables $y_t = p_{1,t}$ and $z_t = m_{1t}$. Conditional on the relative preset price and the policy instrument, consumption is determined by (22); the relative flexible price is determined by (14); allocational efficiency is determined by (18); and marginal cost is determined by (13) and (22). The nominal interest rate is determined residually from equation (17). The policymaker's objective function is

$$E_0 \sum_{t=0}^{\infty} \beta^t \left\{ \ln \left(m_{1t} p_{1,t} \right) - \chi d \left(p_{1,t} \right) m_{1t} p_{1,t} / \xi_t \right\}, \tag{31}$$

and the FOC for profit maximization (15) corresponds to the dynamic constraint (23) for $t \geq 0$:

$$\begin{aligned} 0 \quad = \quad & p_0 \left(p_{1,t} \right)^{1-\varepsilon} \left(1 - \mu \chi \frac{p_{1,t}}{p_0 \left(p_{1,t} \right)} \frac{m_{1t}}{\xi_t} \right) \\ & + \beta E_t \left[p_{1,t+1}^{1-\varepsilon} \left(1 - \mu \chi \frac{m_{1t+1}}{\xi_{t+1}} \right) \right]. \end{aligned} \tag{32}$$

Optimal Policy with Full Commitment

Under full commitment, the policymaker maximizes the value function (31) subject to the constraints (32). The FOCs corresponding to equations (25) for $t > 0$ are

$$0 \quad = \quad \frac{1}{m_{1t}/\xi_t} - \chi d_t p_{1t} - \mu \chi \lambda_t p_{ot}^{-\varepsilon} p_{1t} - \mu \chi \lambda_{t-1} p_{1t}^{1-\varepsilon}, \text{ and} \tag{33}$$

$$0 \quad = \quad \frac{1}{p_{1t}} - \chi \frac{m_{1t}}{\xi_t} d_t - \chi \frac{m_{1t}}{\xi_t} p_{1t} \frac{\partial d_t}{\partial p_{1t}} \tag{34}$$

[3] Another common approach to the analysis of optimal monetary policy starts with a linear-quadratic approximation of the planning problem, e.g., Giannoni and Woodford (2002a, 2002b). For this alternative approach, one obtains a quadratic approximation of the objective function and a linear approximation of the constraints around the steady state of the planning problem and then solves the linear-quadratic (LQ) optimization problem. In general, the results from the two approaches will differ since the LQ approach does not use the second-order terms in the constraint functions, whereas the approach that linearizes the first-order conditions does use this information. Recently, Benigno and Woodford (2005) have shown how to modify the LQ problem such that the analysis of the LQ problem is equivalent to the analysis of the linearized FOCs.

$$+\lambda_t p_{0t}^{-\varepsilon}\left\{(1-\varepsilon)\left(1-\mu\chi\frac{m_{1t}}{\xi_t}\frac{p_{1t}}{p_{0t}}\right)\frac{\partial p_{0t}}{\partial p_{1t}}-\mu\chi\frac{m_{1t}}{\xi_t}\left(1-\frac{p_{1t}}{p_{0t}}\frac{\partial p_{0t}}{\partial p_{1t}}\right)\right\}$$
$$+\lambda_{t-1}p_{1t}^{-\varepsilon}(1-\varepsilon)\left(1-\mu\chi\frac{m_{1t}}{\xi_t}\right).$$

Equation (33) denotes the FOC with respect to real balances, m_1, and equation (34) denotes the FOC with respect to the relative price, p_1.

The Deterministic Steady State of the Full-Commitment Policy

In the deterministic steady state of the full-commitment policy, there is zero inflation (King and Wolman 1999; Wolman 2001). We can easily verify that $\pi = p_0 = p_1 = d = 1$ is indeed a deterministic steady state of equations (32), (33), and (34). Combining equation (13) with the monetary policy equation (22) yields an expression that relates marginal cost to real balances and the preset relative price

$$\psi = \chi m_1 p_1. \tag{35}$$

We can substitute this expression for marginal cost in the FOC for profit maximization of price-adjusting firms, (32), and, using the definition of the inflation rate π, (16), obtain

$$m_1 = \frac{1}{\chi\mu}\frac{1+\beta\pi^{\varepsilon-1}}{1/\pi+\beta\pi^{\varepsilon-1}}. \tag{36}$$

Thus conditional on no inflation, $\pi^{FC} = 1$, real balances are $m_1^{FC} = 1/(\chi\mu)$, and marginal cost is $\psi^{FC} = 1/\mu$. Substituting for marginal cost in the FOC for real balances (33) yields the steady state value for the Lagrange multiplier $\lambda^{FC} = (1-1/\mu)/2$, and the FOC for preset relative prices, (34), is satisfied.

Local Properties of the Full-Commitment Solution

First, we show that the solution to the full-commitment problem stabilizes the prices in response to productivity shocks (King and Wolman 1999). Second, we show that the full-commitment policy rule implements the competitive equilibrium. In the following, let a hat denote the percentage deviation of a variable from its steady state value.

The log-linear approximation of equations (32), (33) and (34) around the no-inflation steady state for $t > 0$ are

$$0 = 2\hat{p}_{1t}+\left(\hat{m}_{1t}-\hat{\xi}_t\right)+\beta E_t\left[\hat{m}_{1,t+1}-\hat{\xi}_{t+1}\right], \tag{37}$$

$$0 = \hat{p}_{1t}+\left(\hat{m}_{1t}-\hat{\xi}_t\right)+\lambda^{FC}\left(\hat{\lambda}_t+\hat{\lambda}_{t-1}\right), \text{ and} \tag{38}$$

$$0 = \left[\mu\frac{2\mu-1}{\mu-1}\right]\hat{p}_{1t}+[1+\chi(\mu-1)]\left(\hat{m}_{1t}-\hat{\xi}_t\right)+(\mu-1)\hat{\lambda}_t. \tag{39}$$

We solve this linear difference equation system through the method of undetermined coefficients. Given the structure of the equation system, it is reasonable to guess that the only relevant state variable is the lagged Lagrange multiplier, λ_{t-1}, and that the solution is of the form

$$\hat{m}_{1t} - \hat{\xi}_t = \gamma \hat{\lambda}_{t-1},\ \hat{p}_{1t} = \theta \hat{\lambda}_{t-1}, \text{ and } \hat{\lambda}_t = \omega \hat{\lambda}_{t-1} \text{ for } t > 0. \tag{40}$$

Now substitute these expressions in equations (37)–(39) and confirm that they solve the difference equation system. This procedure yields three equations that can be solved for the unknowns (ω, γ, ρ).

The optimal full-commitment policy increases normalized real balances m_1 with productivity shocks such that relative prices are not affected, (40). Relative prices respond to past commitments of the policymaker as reflected in the Lagrange multiplier λ, and the Lagrange multiplier evolves independently of productivity shocks. When the Lagrange multiplier attains its steady state value it stays there and optimal policy from thereon fixes the price level and relative prices. We do not prove it, but for reasonable numerical values of (β, μ, χ) the coefficient ω is negative but less than one in absolute value, that is, the system oscillates, but it is stable. In Figure 1 we graph the transitional dynamics of the economy for some parameter values that are standard for quantitative economic analysis, $\beta = 0.99$, $\mu = 1.1$, and $\chi = 1$. As we can see, all variables display dampened oscillations around their steady state values. As discussed above, the FOCs for the initial period of the full-commitment problem are equivalent to the FOCs (38) and (39) with $\lambda_{-1} = 0$, that is, $\hat{\lambda}_{-1} = -1$. Thus during a transition period, as the Lagrange multiplier converges to its steady state value, relative prices change in proportion to the value of the Lagrange multiplier.

The money-supply policy rule, defined as the first and third expression in (40), implements the optimal allocation as a competitive equilibrium. To see this, substitute the policy rule into the log-linear approximation of the optimal pricing equation (37), and we get

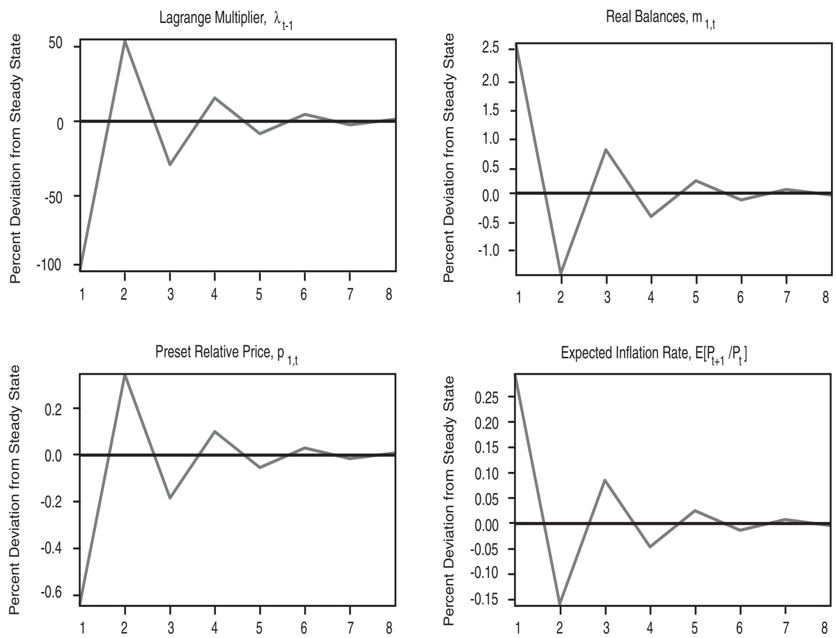
$$\hat{p}_{1t} = \frac{1}{2}\gamma \left(1 + \beta\omega\right) \hat{\lambda}_{t-1}. \tag{41}$$

Thus, conditional on the full-commitment optimal policy rule for real balances, there exists a unique rational expectations equilibrium (REE) for the economy.

Markov-Perfect Optimal Policy

For a Markov-perfect optimal monetary policy, the policymaker at time t maximizes the value function (31) subject to the constraints (32), assuming that future policy choices are some function of the future exogenous shock. The FOCs for this problem correspond to equations (26) for $t = 0$ and are

Figure 1



$$0 = \frac{1}{m_{1t}/\xi_t} - \chi d_t p_{1t} - \mu \chi \lambda_t p_{ot}^{-\varepsilon} p_{1t}, \text{ and} \tag{42}$$

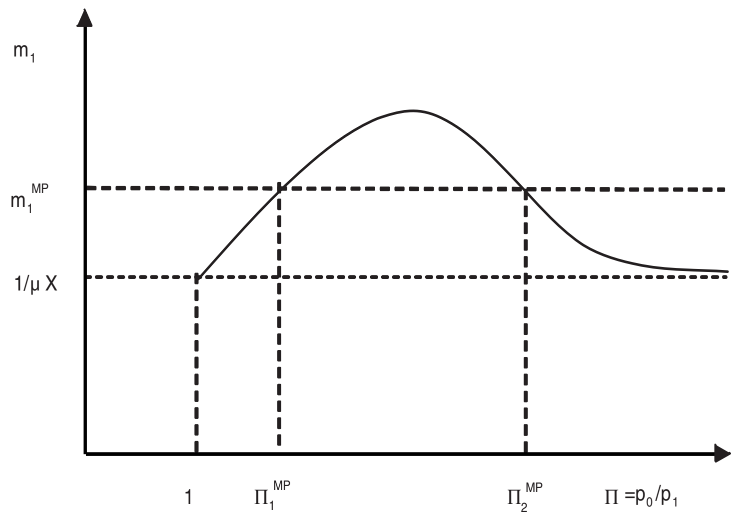
$$\begin{aligned} 0 = & \frac{1}{p_{1t}} - \chi \frac{m_{1t}}{\xi_t} d_t - \chi \frac{m_{1t}}{\xi_t} p_{1t} \frac{\partial d_t}{\partial p_{1t}} \\ & + \lambda_t p_{0t}^{-\varepsilon} \left\{ (1-\varepsilon) \left(1 - \mu \chi \frac{m_{1t}}{\xi_t} \frac{p_{1t}}{p_{0t}} \right) \frac{\partial p_{0t}}{\partial p_{1t}} - \mu \chi \frac{m_{1t}}{\xi_t} \left(1 - \frac{p_{1t}}{p_{0t}} \frac{\partial p_{0t}}{\partial p_{1t}} \right) \right\}. \end{aligned} \tag{43}$$

Equation (42) denotes the FOC with respect to real balances, m_1, and equation (43) denotes the FOC with respect to the relative price, p_1.

The Deterministic Steady State of the Markov-Perfect Policy

The deterministic steady state of the Markov-perfect equilibrium has positive inflation, as opposed to the steady state of the full-commitment solution. It is straightforward to show that optimal policy does not stabilize prices in the steady state. Suppose to the contrary that there is no inflation in the steady state, $p_0 = p_1 = 1$, then evaluating equations (32), (42), and (43) at their deterministic steady state implies that $\partial d/\partial p_1 < 0$. But with stable prices,

Figure 2



$\pi = 1$, the derivative of allocational efficiency with respect to p_1,

$$\partial d/\partial p_1 = \varepsilon p_1^{-\varepsilon-1}\left(\pi^{-\varepsilon-1} - 1\right), \tag{44}$$

is zero, and we have a contradiction. On the other hand, with positive inflation, the impact of p_1 on allocational efficiency is negative. This suggests that the steady state inflation rate is positive, as indeed shown by Wolman (2001). We can find the steady state inflation rate as the solution to the following fix-point problem. Conditional on some inflation rate, π, use equations (35) and (36) to determine steady state real balances, m_1, and marginal cost, ψ. Conditional on (π, m_1, ψ), use equation (42) to obtain the steady state Lagrange multiplier λ. Finally, we have to verify that equation (43) is satisfied.

The competitive equilibrium constraint (32), together with the FOCs for optimal policy, (42) and (43), evaluated at their deterministic steady state indeed yield a unique solution for the steady state, $\left(\pi^{MP}, m_1^{MP}, \psi^{MP}\right)$. Note, however, that contingent on the steady state Markov-perfect real balances m_1^{MP}, the competitive equilibrium constraint alone is consistent with multiple steady states. In Figure 2, we graph real balances as a function of the inflation rate, π, based on equation (36). Notice that as the inflation rate increases, real balances first increase and then decline. This means that for a given choice of real balances that is not too high, $m_1 > m_1^{FC} = 1/(\chi\mu)$, there are two steady state inflation rates.

Local Properties of the Markov-Perfect Policy

For a local approximation of the optimal Markov-perfect policy we can show that the policy stabilizes prices around the trend growth path in response to productivity shocks. Because the steady state involves positive inflation, the expressions for the local approximations are quite convoluted, and we do not display them here. Suffice it to say that locally the optimal Markov-perfect solution is of the form

$$\hat{p}_{1t} = \hat{m}_{1t} - \hat{\xi}_t = \hat{\lambda}_t = 0. \tag{45}$$

We can substitute the local approximation of the Markov-perfect policy rule, second and third equalities of (45), into the log-linear approximation of the optimal pricing equation (15) when the steady state has non-zero inflation and get

$$\hat{p}_{1t} = \left[\beta \frac{(\varepsilon - 1)(1 - \mu\chi m_1)\pi^{\varepsilon}}{(\varepsilon - 1)\pi^{e} - \mu\chi m_1 \left(1 + \varepsilon\pi^{\varepsilon-1}\right)} \right] \hat{p}_{1,t+1}. \tag{46}$$

Note that for a steady state with zero inflation, the coefficient on the right-hand side term is zero. Since the steady state of the Markov-perfect equilibrium involves only a very small amount of inflation, the coefficient on future prices is close to zero and certainly less than one. Thus, solving the equation forward implies that there exists a unique REE, $\hat{p}_{1t} = 0$.

4. GLOBAL PROPERTIES OF OPTIMAL POLICY

We now show that the policy rule implied by a Markov-perfect optimal policy does not globally implement the optimal policy allocation. We also conjecture that the policy rule implied by the full-commitment policy may not always be implementable. An augmented full-commitment policy rule that can respond to contemporaneous variables as described in Section 2, however, is likely to implement the optimal policy allocation.

For the analysis of the global properties of policy rules, it will be useful to rewrite a firm's profit maximization condition (12), which represents the competitive equilibrium constraint for the planning problem. Solve this expression for a firm's optimal relative price as a markup over the average marginal cost for which the price is set

$$\frac{P_t(j)}{P_t} = \mu \frac{\psi_t + \beta E_t \left[\psi_{t+1} (P_{t+1}/P_t)^{\varepsilon}\right]}{1 + \beta E_t \left[(P_{t+1}/P_t)^{\varepsilon-1}\right]}. \tag{47}$$

We can think of this expression as a firm's optimal relative price choice on the left-hand side, p_{0t}, conditional on the relative prices set by all other firms, $\bar{p}_{0t}$, determining the right-hand side of the equation. The behavior of the other firms is reflected in the equilibrium values of marginal cost and the inflation rate. For our argument, we will assume that there are no shocks to the economy, that is, productivity is constant. Using the equilibrium conditions (13), (16), and (22) for the right-hand side of (47), we then get

$$p_{0,t} = \mu\chi \frac{m_{1t} p_1\left(\bar{p}_{0,t}\right) + \beta m_{1,t+1} \bar{p}_{0,t} \pi_{t+1}^{\varepsilon-1}}{1 + \beta \pi_{t+1}^{\varepsilon-1}} \text{ with } \pi_{t+1} = \bar{p}_{0,t} / p_1\left(\bar{p}_{0,t+1}\right). \tag{48}$$

Markov-Perfect Policy

The Markov-perfect policy rule not only stabilizes prices in response to small productivity shocks, but stabilization is the globally optimal response to shocks,

$$m_{1t} = m_1^{MP} \xi_t. \tag{49}$$

We can verify that (49) is the optimal response to productivity shocks by substituting the expression for m_{1t} into equations (32), (42), and (43). This policy rule reflects the definition of a Markov-perfect policy: it depends only on payoff-relevant state variables, that is, ξ_t only in our case.

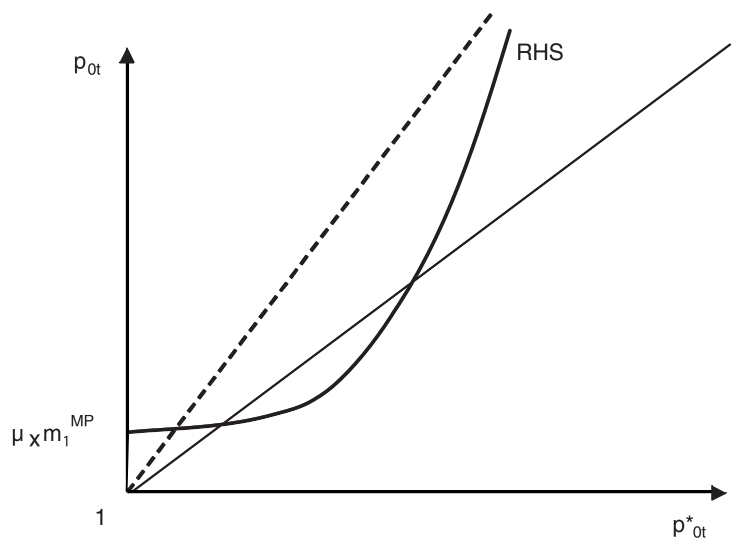
In general, the Markov-perfect policy rule cannot implement the planning allocation as a competitive equilibrium outcome. King and Wolman (2004) argue that a Markov-perfect optimal policy introduces strategic complementarities into the firms' price-setting behavior and thereby makes multiple equilibria possible. With constant normalized real balances of the Markov-perfect policy and no productivity shocks, the optimal pricing condition (48) simplifies to

$$p_{0,t} = \mu\chi m_1^{MP} \frac{p_1\left(\bar{p}_{0,t}\right) + \beta \bar{p}_{0,t} \pi_{t+1}^{\varepsilon-1}}{1 + \beta \pi_{t+1}^{\varepsilon-1}}. \tag{50}$$

Strategic complementarities are said to be present if a representative firm increases its own control variable when it perceives that all other firms increase their control variable. In terms of the price-setting equation (50): a firm increases its own relative price, p_{0t}, on the left-hand side of the expression if all other firms increase their relative price, $\bar{p}_{0t}$, on the right-hand side of the expression. Essentially, if all other firms increase their price, $\bar{p}_{0t}$, then the expected inflation rate increases, and therefore a firm will increase its own relative price in order to prevent an erosion of its relative price in the next period. Since the equilibrium relative price is a fix-point of expression (50), $p_{0t} = \bar{p}_{0t}$, strategic complementarities raise the possibility of multiple fixed points, that is, multiple equilibria.

In Figure 3 we graph the RHS of (50), conditional on some value for $p_{1,t+1}$. If we evaluate the RHS of (50) at $\bar{p}_{0t} = 1$, we get $p_{0t} = 1$ and $RHS = \mu\chi m_1^{MP} > 1$. If we consider the limit of the RHS as $\bar{p}_{0t}$ becomes arbitrarily large, we see that $\bar{p}_{1t}$ converges to a finite value and the inflation rate becomes arbitrarily large, thus the RHS converges to a line through the

Figure 3



origin with slope $\mu\chi m_1^{MP} > 1$. Without a further analysis of the behavior of the RHS for finite positive values of $\bar{p}_{0t}$, this at least suggests the possibility of two intersection points of the RHS with the diagonal. Furthermore we know that in the steady state, when $p_{1,t+1} = p_1^{MP}$, there are indeed two solutions for p_0 to equation (36). King and Wolman (2004) show that, in general, there exist two intersection points. Thus there is no unique equilibrium and the Markov-perfect policy rule does not implement the planning allocation.

Full-Commitment Policy

Optimal full-commitment monetary policy stabilizes prices in response to productivity shocks not only locally around the steady state, but also globally,

$$m_{1t} = \tilde{m}_{1t}\xi_t, \tilde{m}_1 = \Gamma(\lambda_{t-1}), p_{1t} = \Theta(\lambda_{t-1}), \text{ and } \lambda_t = \Omega(\lambda_{t-1}). \quad (51)$$

To see this, simply note that equations (32), (33), and (34) define a system in $(\tilde{m}_{1t}, p_{1t}, \lambda_{t-1})$ that is independent of productivity shocks. Different from the Markov-perfect policy, the Lagrange multiplier on the competitive equilibrium constraint is not constant and therefore the normalized real balances are not constant.

We do not have unambiguous results on the implementation of the planning allocation through the full-commitment policy rule. On the one hand, we can show that if the Lagrange multiplier has attained its steady state value, $\lambda_{t-1} = \lambda^{FC}$, then the full-commitment policy rule implements the planning solution. On the other hand, as long as the Lagrange multiplier has not attained its steady state, the full-commitment policy rule suffers from some of the same problems as does the Markov-perfect policy.

Suppose that the Lagrange multiplier has attained its steady state value, $\lambda_{t-1} = \lambda^{FC}$. If we substitute the value for the Lagrange multiplier in the FOCs (33) and (34), we can see that they will always be satisfied from there on. But this means that from there on the normalized real balances attain their steady state value, m_1^{FC}, and the competitive equilibrium constraint (32) simplifies to

$$0 = p_{0t}^{1-\varepsilon}\left(1 - \frac{p_{1t}}{p_{0t}}\right). \tag{52}$$

Therefore $p_{1t} = p_{0t}$, that is $P_t^{FC} = P_{t-1}^{FC}$, and prices are determined.

Now consider the transitional phase when the Lagrange multiplier differs from its steady state value. Given the implied policy rule (51), we can construct future nominal money stocks recursively as functions of the initial value of the Lagrange multiplier

$$M_t = \xi_t \cdot \Gamma(\lambda_{t-1}) \cdot \Theta(\lambda_{t-1}) \cdot P_t, \text{ and} \tag{53}$$

$$P_t = \frac{p_{0,t-1}}{p_{1,t}} P_{t-1} = \frac{p_0\left(p_{1,t-1}\right)}{p_{1,t}} P_{t-1} = \frac{p_0\left[\Theta(\lambda_{t-2})\right]}{\Theta(\lambda_{t-1})} P_{t-1}. \tag{54}$$

With full commitment, a policymaker can always announce a time path for the nominal money supply and follow through on that announcement. Given the nominal money supply rule, we can rewrite the optimal pricing condition (48) in nominal terms and get

$$P_{0t} = \mu\chi \frac{M_t + \beta M_{t+1}\left(P_{t+1}/P_t\right)^{\varepsilon-1}}{1 + \beta\left(P_{t+1}/P_t\right)^{\varepsilon-1}} \tag{55}$$

$$\text{with } \frac{P_{t+1}}{P_t} = \left[\frac{P_{0,t+1}^{1-\varepsilon} + P_{0,t}^{1-\varepsilon}}{P_{0,t}^{1-\varepsilon} + P_{0,t-1}^{1-\varepsilon}}\right]^{1/(1-\varepsilon)}.$$

As we do for the analysis of the Markov-perfect policy, we are looking for a fix point in the optimal nominal price, P_{0t}, conditional on the past and future nominal prices, $P_{0,t-1}$ and $P_{0,t+1}$, and the nominal money stocks, M_t and M_{t+1}. Clearly for a constant money supply, that is, the constant steady state Lagrange multiplier, there is a unique solution for P_{0t}. If the Lagrange multiplier converges globally to its steady state, then if the difference between M_t and M_{t+1} is small enough, we will also have a unique solution. We do not, however, prove that there is a unique solution for the initial phase of the transition period.

Note that for full-commitment policy, we have only outlined the same potential for multiple equilibria as King and Wolman (2004) have shown to exist for the Markov-perfect policy rule. We have not proven that the full-commitment policy rule cannot implement the planning allocation. Whether or not the full-commitment policy rule implements the planning allocation may be irrelevant if one believes that a policymaker can always respond to contemporaneous variables. If such a response is feasible, then an augmented full-commitment policy rule as described in Section 2 may always implement the planning allocation.

5. CONCLUSION

This paper has considered optimal monetary policy as the solution to both full-commitment and time-consistent Markov-perfect planning problems. The solutions are consistent with rational expectations competitive equilibria. The optimal solution to the planning problem implies a rule for the assumed policy instrument, in our case, a money supply instrument. We have then verified that, for local approximations to the solution of the optimal policy problem, the implied policy rules implement the planning allocations, that is, the planning allocation is the unique rational expectations equilibrium conditional on the implied policy rule. However, following on the insights of King and Wolman (2004), we have then examined whether the implied policy rules also implement the allocation globally. We find that a money supply rule that is Markov-perfect does not implement the planning solution. We provide a partial argument that the full-commitment money supply rule does implement the planning solution, but we do not have a complete proof for this statement.

For the analysis, we have taken the choice of monetary instrument, in this case the nominal money stock, as given but this choice is not innocuous. In other work (Dotsey and Hornstein 2005), we have argued that equilibrium indeterminacy may depend on the choice of policy instrument. In particular, if the Markov-perfect policy uses the nominal interest rate as an instrument, the equilibrium is determinate.

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Can Feedback from the Jumbo CD Market Improve Bank Surveillance?

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In recent years, policymakers in the Basel countries have begun exploring strategies for harnessing financial markets to contain bank risk. Indeed, the new Accord counts market discipline, along with supervisory review and capital requirements, as an explicit pillar of bank supervision.[1] A popular proposal for implementing market discipline in the United States would require large banks to issue a standardized form of subordinated debt (Board of Governors 1999; Board of Governors 2000; Meyer 2001). Advocates of

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[1] The cornerstone of supervisory review—the most important of the pillars—is thorough, regularly scheduled, on-site examinations. The Federal Deposit Insurance Corporation Improvement Act of 1991 (FDICIA) requires most U.S. banks to submit to a full-scope examination every 12 months. These examinations focus on six components of safety and soundness—capital protection (C), asset quality (A), management competence (M), earnings strength (E), liquidity risk exposure (L), and market risk sensitivity (S)—CAMELS. At the close of each exam, an integer ranging from 1 (best) through 5 (worst) is awarded for each component. Supervisors then use these component ratings to assign a composite CAMELS rating reflecting overall condition—also on a 1-to-5 scale. In general, banks with composite ratings of 1 or 2 are considered satisfactory while banks with ratings of 3, 4, or 5 are unsatisfactory and subject to supervisory sanctions. (Footnote 10 offers more details about these sanctions.) At year-end 2005, 4.63 percent of U.S. banks held unsatisfactory ratings.

this proposal argue that high-powered performance incentives in the subordinated debt (sub-debt) market will produce accurate risk assessments. And, in turn, these assessments—expressed for risky institutions through rising yields or difficulties rolling over maturing debt—will pressure bank managers to maintain safety and soundness (Calomiris 1999; Lang and Robertson 2002).

Even if financial markets apply little direct pressure to curb risk taking, market data could still enhance supervisory review by improving off-site surveillance.[2] Off-site surveillance involves the use of accounting data and anecdotal evidence to monitor the condition of supervised institutions between scheduled exams.[3] Market assessments could enhance surveillance in three ways: (1) by flagging banks missed by conventional off-site tools, (2) by reducing uncertainty about banks flagged by other tools, or (3) by providing earlier warning about developing problems in banks flagged by these tools (Flannery 2001). Such enhancements would reduce failures over time by enabling supervisors to take action earlier to address safety-and-soundness problems.

One concern about attempts to incorporate market data into surveillance is regulatory burden—current proposals would require large banking organizations to float a standardized issue of sub-debt. That most large banks currently issue sub-debt does not imply the burden is negligible.[4] Voluntary issuance varies considerably over time with market conditions. For example, the number of sub-debt issues by the top-50 banking organizations rose from 3 in 1988 to 108 in 1995, only to fall to 42 in 1999 (Covitz, Hancock, and Kwast 2002). Moreover, banks currently issuing sub-debt may be choosing maturities unlikely to produce valuable risk signals, so a mandated maturity would still impose a regulatory burden. Before placing additional burden on the banking sector, particularly at a time when other sizable regulatory changes (Basel II) are in the offing, supervisors should first assess the power of risk signals from existing securities.

One potential source of risk assessments that can be mined without increasing regulatory burden is the market for jumbo certificates of deposit (CDs).

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[2] Bliss and Flannery (2001) found that managers of holding companies do not respond to market pressure to contain risk, though Rajan (2001) questioned the ability of their framework to unearth such evidence.

[3] Examination is the most effective tool for spotting safety-and-soundness problems, but it is costly and burdensome—costly because of the examiner resources required and burdensome because of the intrusion into bank operations. Surveillance reduces the need for unscheduled visits by prodding bankers to contain risk between scheduled exams. It also helps supervisors plan exams by highlighting risk exposures. For example, if pre-exam surveillance reports indicate a bank has significant exposure to interest rate fluctuations, supervisors will staff the exam team with additional market risk expertise.

[4] Mandating issuance of a security with specific attributes is tantamount to a tax on capital structure. Although we know of no direct evidence about the burden of this tax, heterogeneity in sub-debt maturities, outstanding volume over time, and the source of issue (bank vs. bank holding company) suggest it is nontrivial.

Jumbo CDs are time deposits with balances exceeding $100,000. The typical bank relies on a mix of deposits to fund assets—checkable deposits, passbook savings accounts, retail CDs, and jumbo CDs. Both retail and jumbo CDs have fixed maturities (as opposed to checkable deposits which are payable on demand); they differ by Federal Deposit Insurance Corporation (FDIC) coverage. Only the first $100,000 of deposits is eligible for insurance, so the entire retail CD (which is less than $100,000) is insured while only the first $100,000 of a jumbo CD is covered. Checkable deposits, passbook savings, and retail CDs are often collectively referred to as "core deposits" because balances respond little to changes in bank condition and market rates. Full FDIC coverage makes these deposits a stable and cheap source of funding. At year-end 2005, U.S. banks funded on average 67.1 percent of assets with core deposits and 14.4 percent with jumbo CDs. The average jumbo CD balance in the fourth quarter of 2005 was $330,886; the average balance in 95 percent of the U.S. banks exceeded $152,115. The average maturity was just over one year. Jumbo CDs are considered a "volatile" liability because relatively large uninsured balances and short maturities force issuing banks to match yields (risk-free rates plus default premiums) available in the money market or lose the funding. This pressure to "price" new conditions quickly makes the jumbo CD market, in theory, an important source of feedback for off-site surveillance.[5]

Potentially valuable jumbo CD data are currently available for most commercial banks. In contrast, only very large banking organizations now issue sub-debt. These organizations may be the most important from a systemic-risk standpoint, but the focus of off-site surveillance—indeed of all U.S. prudential supervision—is on the bank, and most banks do not issue or belong to holding companies that issue sub-debt. Moreover, a negative risk signal from a holding company claim would not, by itself, help supervisors identify the troubled subsidiary. Jumbo CDs constitute a large class of direct claims on both large and small banks. At year-end 2005, U.S. banks with more than

[5] Since the early 1990s, financial innovation has offered households a growing array of substitutes for traditional bank deposits. As a result, the supply of core deposits has declined secularly, forcing banks to turn to more volatile funding sources such as jumbo CDs. Between 1992 and 2005, for example, the average core deposit-to-total asset ratio for U.S. banks tumbled from 80.1 percent to 67.1 percent, while average jumbo CD dependence jumped from 7.5 percent to 14.4 percent of assets. Increasing reliance on jumbo CDs implies greater exposure to liquidity and market risk—a bad outcome from the perspective of a bank supervisor. At the same time, the $100,000 ceiling on deposit insurance makes jumbo CD holders savvier about bank risk than other depositors. So the jumbo CD market could exert pressure on bank managers to contain risk—either directly through the impact of higher yields and lower balances on profits or indirectly through supervisory responses to risk signals conveyed by yields and withdrawals. Such pressure would complement supervisory review. Hence, another contribution of this article is to offer insight into the tradeoff by quantifying the potential contribution of jumbo CD data to off-site surveillance. See Feldman and Schmidt (1991) for further discussion of the tradeoff between greater risk exposure and more reliable market data implied by rising jumbo CD dependence. Our results suggest this rising dependence makes supervisors on balance worse off.

$500 million in assets funded 14.6 percent of assets with jumbo CDs; for banks with less than $500 million, the average jumbo-CD-to-total-asset ratio was 14.3 percent. Finally, risk signals in the form of yields and withdrawals can be cheaply and easily constructed because banks report jumbo CD interest expense and balances quarterly to their principal supervisor. Also, nearly 30 years of research—much of which relies on these interest-expense and account-balance data—has produced robust evidence of risk pricing in the jumbo CD market.

Data from the jumbo CD market might prove particularly useful in community bank surveillance. Community banks specialize in making loans to and taking deposits from small towns or city suburbs. For regulatory purposes, the Financial Modernization Act of 1999 established an asset threshold of $500 million—expressed in constant 1999 dollars. At year-end 2005, nearly 90 percent of U.S. banks operated on this scale. Not surprisingly, most failures are community banks. They also frequently operate on extended exam schedules, with up to 18 months elapsing between full-scope, on-site visits. This schedule diminishes the quality of quarterly financial statements, thereby reducing the effectiveness of off-site monitoring.[6] It is possible that holders of community bank jumbo CDs supplement public financial data with independent "Peter-Lynch-type" research.[7] Or, inside information about bank condition could leak from boards of directors, which typically include prominent local businesspeople. (Community bank jumbo CDs are often held by such "insiders.") Thus, sudden changes in yields or withdrawals might signal trouble more quickly or reliably than surveillance tools based on financial statements.

In short, jumbo CDs fund a large portion of bank assets and furnish a cheap source of market data, yet no study has formally tested the surveillance value of yields and withdrawals. We do so with an early warning model and out-of-sample timing conventions designed to mimic current surveillance practices. Specifically, we generate risk rankings using jumbo CD default premiums and quarter-over-quarter withdrawals for banks with satisfactory supervisory ratings. We rank the same banks by CAMELS-downgrade probability as estimated by an econometric surveillance model. Finally, out-of-sample performance for all three rankings is compared over a sequence of two-year windows running from 1992 to 2005, counterfactually as if super-

[6] Verification of financials is an important source of value created by exams (Berger and Davies 1998; Flannery and Houston 1999). Indeed, recent research has documented large adjustments in asset-quality measures following on-site visits, particularly for banks with emerging problems (Gunther and Moore 2000).

[7] Peter Lynch ran Fidelity's Magellan Fund from 1977 to 1990. During this period, fund value rose over 2,700 percent. Lynch was famous for looking past financial statements to the real world, observing consumer and firm behavior in malls, for example. For more details, see Lynch and Rothchild (2000).

visors in the fourth quarter of each year possessed data only up to that point. We find that jumbo CD signals would not have flagged banks missed by the CAMELS-downgrade model or would not have reduced uncertainty about banks flagged by the model. We also find that jumbo CD signals would not have provided earlier warning about developing problems in banks flagged by the CAMELS-downgrade model. These results are broadly consistent with other recent work, so we close by exploring reasons the surveillance value of market data may have been overestimated.

1. PRIOR LITERATURE

Research on the jumbo CD market since the mid-1970s—mostly with 1980s data—has consistently found evidence of risk pricing (see Table 1). Some 20 articles have been published using a mix of time series and panel approaches: 18 articles exploited U.S. data, 11 examined only yields, 4 examined only runoff (i.e., deposit withdrawals), and 5 studied both. Most drew heavily on quarterly financial statements. Only one article—the first contribution to the literature in 1976—found no link between bank risk and yields or runoff. In some ways, the robustness of these results is striking because U.S. samples mostly predate the Federal Deposit Insurance Corporation Improvement Act of 1991 (FDICIA). Before this Act, the majority of failures were resolved through purchases and assumptions, whereby the FDIC offered cash to healthy banks to assume the liabilities of failed ones. So, even though jumbo CD holders faced default risk in theory, many were shielded from losses in practice.[8]

Although evidence from prior literature about our out-of-sample test windows (1992–2005) is thinner, intuition and history make a case for significant risk sensitivity. The handful of articles looking at 1990s data found risk pricing, but no study examined jumbo CD data for the post-2000 period. Nonetheless, economic intuition suggests sensitivity should be strong because of three important institutional changes in the 1990s. First, as noted, the FDICIA directed the FDIC to resolve failures in the least costly way, which implies imposing a greater share of losses on uninsured bank creditors (Benston and Kaufman 1998; Kroszner and Strahan 2001).[9] This change should

[8] Before 1991, expected losses had three components: (1) the probability of bank failure, (2) the loss if the failed bank were not purchased by a healthy one, and (3) the probability the failed bank would not be purchased. Even if (1) and (2) were positive, expected losses would still be approximately zero if jumbo CD holders expected all failures to be resolved with purchase and assumptions. The need to model FDIC behavior, therefore, complicates estimation of risk sensitivity for the pre-1991 regime. Suppose, for example, (1) and (2) fall, reducing expected losses, incentives to monitor risk, and jumbo CD risk sensitivity. But the FDIC responds by curtailing implicit coverage—perhaps because of the reduced threat of contagious runs. If large enough, this offsetting effect could induce a rise in measured sensitivity to bank condition.

[9] As discussed in footnote 8, expected losses equal zero if jumbo CD holders anticipate resolution through purchase and assumptions. But the FDICIA should have changed expectations

have increased expected losses for jumbo CD holders and their incentive to monitor bank condition. Second, the Financial Institutions Reform, Recovery, and Enforcement Act of 1989 required supervisors to disclose serious enforcement actions (Gilbert and Vaughan 2001).[10] Third, in the late 1990s, the FDIC began putting quarterly financial data for individual banks on the Web, along with tools for comparing performance with industry peers. The second and third change should have lowered the cost to jumbo CD holders of monitoring bank condition. Evidence from U.S. banking history also implies our sample should feature strong risk pricing. Gorton (1996), for example, documented a link between discounts on state bank notes and issuer condition during the free-banking era, while Calomiris and Mason (1997) observed sizable differences in yields and runoff for weak and strong Chicago banks prior to the 1932 citywide panic. Friedman and Schwartz (1960) also noted that public identification of banks receiving loans from the Reconstruction Finance Corporation triggered runs in August 1932. More recently, Continental Illinois began hemorrhaging uninsured deposits when the extent of its problems became public in May 1984 (Davison 1997). In all these cases, uninsured claimants monitored and reacted to changes in bank condition, thereby impounding risk assessments into prices or quantities.

Evidence of risk pricing in the jumbo CD market does not imply that yield and runoff data would add value in surveillance. First, stable in-sample estimates of reactions to current bank condition and reliable out-of-sample forecasts of emerging safety-and-soundness problems are not the same thing. Evidence from the market efficiency literature, for example, has demonstrated that trading strategies based on well-documented pricing anomalies, such as calendar effects, size effects, and mean revision, do not offer abnormal returns when tested in real time by fund managers (Roll 1994; Malkiel 2003). Second, just as assessing the profitability of trading rules requires a benchmark, such as the return from an index fund, assessing the surveillance value of market data requires a baseline for current practices. It is not enough to note that jumbo CD signals flag problem banks because supervisors already have systems in place for these purposes. The true litmus test is this: Does integration of

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about FDIC behavior. Between 1988 and 1990, jumbo CD holders suffered losses in only 15 percent of bank failures. From 1993 to 1995, they lost money 82 percent of the time.

[10] The term "enforcement action" refers to a broad range of powers used to address suspect practices of depository institutions and institution-affiliated parties—the supervisory sanctions mentioned in footnote 1. Typically, these actions are imposed in response to adverse exam findings, but they can also be triggered by deficient capital levels under Prompt Corrective Action or by negative information gathered through off-site surveillance. Usually enforcement actions are implemented in a graduated manner, with informal preceding formal actions. An informal action is the most common; it is simply a private, mutual understanding between a bank and its supervisory agency about the steps needed to correct problems. Formal actions are far more serious. Supervisors resort to them only when violations of law or regulations continue or when unsafe and abusive practices occur. Formal enforcement actions are legally enforceable and, in most cases, publicly disclosed.

yields and runoff into actual surveillance routines consistently and materially improve out-of-sample forecast accuracy?[11]

Four recent articles have gauged the surveillance value of market data against a current practices benchmark. Evanoff and Wall (2001) compared regulatory capital ratios and sub-debt yields as predictors of supervisory ratings, finding that sub-debt yields modestly outperform capital ratios in one-quarter-ahead tests. Gunther, Levonian, and Moore (2001), meanwhile, observed in-sample improvement in model fit when estimated default frequencies (EDFs, as produced by Moody's KMV) were included in an econometric model designed to predict holding company supervisory ratings with accounting data. Krainer and Lopez (2004) also experimented with equity market variables—in this case, cumulative abnormal stock returns as well as EDFs—in a model of holding company ratings. Unlike Gunther, Levonian, and Moore (2001), they assessed value added in one-quarter-ahead forecasts. Like Evanoff and Wall (2001), they noted only a modest improvement in out-of-sample performance. Finally, Curry, Elmer, and Fissel (2003) added various equity signals to an econometric model built to predict four-quarter-ahead supervisory ratings, again witnessing only a slight increase in forecast accuracy.

Recent tests against a surveillance benchmark have advanced the market data literature, to be sure, but the absence of empirical tests modeled on actual practice mutes the potential impact on supervisory policy. Evanoff and Wall (2001), for example, proxied supervisor perceptions of safety and soundness with regulatory capital ratios—a practice that was problematic because capital is the sole criterion only when Prompt Corrective Action (PCA) thresholds are violated. Otherwise, a variety of measures are weighed.[12] In addition, Gunther, Levonian, and Moore (2001) and Krainer and Lopez (2004) conducted performance tests with holding company data—a problematic approach because, as noted, off-site surveillance focuses on individual banks. Indeed, the Federal Reserve, which has responsibility for holding company supervision, does not maintain an econometric model estimated on holding company data.[13] Finally, Gunther, Levonian, and Moore (2001) and Curry, Elmer, and Fissel (2003) relied on tests unlikely to impress supervisors: the first assessing

[11] As discussed in footnote 10, supervisors use enforcement actions to induce banks to address safety-and-soundness problems. Some are quite severe, going as far as permanent removal from the banking industry. The earlier actions are imposed, the more likely problems can be corrected. But enforcement actions impose significant costs on the bank, so supervisors prefer to wait for compelling evidence of serious problems. Hence, jumbo CD signals could add supervisory value by reinforcing conclusions yielded by other surveillance tools, thereby facilitating swifter action.

[12] Footnote 1 discusses the CAMELS framework supervisors use to assess bank condition. In any event, evidence from counterfactual applications of PCA to late 1980s/early 1990s data (Jones and King 1995; Peek and Rosengren 1997) suggests the thresholds are too low to affect supervisor behavior.

[13] Each article estimated a unique holding company model to benchmark surveillance procedures. Both tested joint hypotheses: (1) the model approximates the one the Fed would use and (2) equity market signals enhance the performance of that model.

in-sample performance only and the second assessing out-of-sample performance with a contemporaneous holdout (rather than a period-ahead sample). Our work improves on this research by employing an econometric model used in surveillance, out-of-sample timing conventions patterned on current practices, and data taken from bank (rather than holding company) financial statements and supervisor assessments. Even more important, we contribute a coherent framework for use in future research on the surveillance value of market data.

2. THE DATA

To test the surveillance value of jumbo CD data, we built a long panel containing financial data and supervisory assessments for all U.S. commercial banks. This data set contained income statement and balance sheet series as well as CAMELS composite and management ratings from 1988:Q1 through 2005:Q4.[14] The accounting data came from the Call Reports—formally the Reports of Condition and Income—which are collected under the auspices of the Federal Financial Institutions Examination Council (FFIEC). The FFIEC requires all U.S. commercial banks to submit such data quarterly to their principal supervisor; most reported items are publicly available. CAMELS ratings were pulled from a nonpublic portion of the National Information Center database; only examiners, analysts, and economists involved in supervision at the state or federal level can access these series. Only one substantive sample restriction was imposed—exclusion of banks with operating histories of under five years. Financial ratios for these start-up, or de novo, banks often take extreme values that do not imply safety-and-soundness problems (DeYoung 1999). For instance, de novos often lose money in their early years, so earnings ratios are poor. Extreme values could introduce considerable noise into risk rankings, making it more difficult to assess relative performance. Another reason for dropping de novos is that supervisors already monitor these banks closely. The Federal Reserve, for example, examines newly chartered banks every six months until they earn a composite rating of 1 or 2 in consecutive exams.

Although our testing framework improves on prior research, our data still contain measurement error. Only a small number of money center banks issue negotiable instruments that are actively traded, so true market yields are not available for a cross section of the industry. It is possible, however, to construct average yields from the Call Reports for all U.S. banks by dividing quarterly in-

[14] Two data notes: (1) Explicit assessment of market risk sensitivity (S) was added in 1997, so pre-1997 composites are CAMEL ratings, and (2) none of our empirical exercises exploits the entire dataset (1988:Q1–2005:Q4); each uses a suitable sub-sample. For example, estimation of the downgrade model ends in 2003 to permit out-of-sample tests on 2004–2005.

terest expense by average balance. Subtracting rates on comparable-maturity Treasuries from these yields produces something that looks like a default premium series. Other researchers have successfully tested hypotheses about bank risk with this approach (for example, James 1988; Keeley 1990; and, more recently, Martinez-Peria and Schmukler 2001). Still, two related types of measurement error must be acknowledged; the proxy is an average rather than marginal measure (and, therefore, somewhat backward looking), and it is a quarterly accounting rather than real-time economic measure.

Measurement error in this series does not imply that jumbo CD data taken from the Call Report lack surveillance value. Jumbo CD holders may react to rising risk by withdrawing funds, and changes in account balances (deposit runoff) can be measured error-free with accounting data.[15] Moreover, distress models based on financial statements have been a cornerstone of public- and private-sector surveillance for decades (Altman and Saunders 1997). Indeed, federal and state supervisors alike give heavy weight to book-value measures of credit risk and capital protection in routine surveillance, yet both contain serious measurement error (Barth, Beaver, and Landsman 1996; Reidhill and O'Keefe 1997). Finally, and most importantly, the supervisory return on jumbo CD signals—or any market signal for that matter—depends not on the value of the signal alone, but rather on that value net of the cost of extraction. Current surveillance routines are built around the Call Reports and, as noted, these reports already contain the data necessary to construct yield and runoff series for jumbo CDs. Even if the marginal surveillance value of jumbo CD signals were low relative to pure market signals because of measurement error, the marginal cost of extracting jumbo CD signals is near zero. The cost of integrating market signals into off-site surveillance is not as low because of the regulatory burden associated with any compulsory security issues and the training burden associated with changes in supervisory practices. It is possible, therefore, that jumbo CD data add more net value than pure market signals. In short, the surveillance value of jumbo CD data is ultimately an empirical issue.

Still, the net contribution of jumbo CD signals to surveillance cannot be positive if measurement error renders the data hopelessly noisy. So, as a check, we performed a simple test on yields and another on runoff—both suggested that bank condition is priced. In the first test, we compared quarterly yields—that is, jumbo CD interest expense divided by average balance—for the 5 percent of banks most and least at risk of failure each year from 1992 to 2005

[15] In the literature, "runoff" is used loosely as a synonym for withdrawals. For this test, we define it as quarter-over-quarter percentage changes in a bank's total dollar volume of jumbo CDs. Later, we define "simple" deposit runoff similarly.

(the period used in out-of-sample testing).[16] Over this period, yields at high-risk banks topped yields at low-risk banks by an average of 25 basis points. (By way of comparison, the average spread between yields on three-month nonfinancial commercial paper and three-month Treasury bills for 1992 to 2005 was 24 basis points.) Institutional changes in the 1990s appear to have strengthened risk pricing. Despite declining money market rates, the mean spread between "risky" and "safe" banks climbed from 14 basis points for 1992–1997 to 33 for 1998–2005 (difference significant at 1 percent). In the second test, we examined quarterly jumbo CD growth at the 169 U.S. banks that failed between 1992 and 2005 for two distinct periods in the migration to failure: two to four years out and zero to two years out. Mean growth two to four years prior to failing was a healthy 8.4 percent. But in the final two years, quarterly growth turned sharply negative, averaging -4.0 percent—a pattern consistent with jumbo CD holders withdrawing funds to avoid losses.

As a final check, we regressed yields and runoff on failure probability and suitable controls; the results also attested to risk pricing. The sample contained observations for all non-de-novo banks with satisfactory supervisory ratings from 1988:Q1 to 2004:Q4.[17] (Table 2 contains the results.) Both coefficients of interest were "correctly" signed and significant at the 1-percent level, implying a rise in failure risk translated into higher yields and larger runoff: coefficient magnitudes were economically small, but it is important to remember that risk sensitivity is a cardinal concept whereas risk ranking is an ordinal one. Recent back-testing of the Focus Report highlights the difference. The Focus Report is a Call-Report-based, Federal Reserve tool for predicting the impact of a 200-basis-point interest rate shock on bank capital. For the 1999–2002 interest rate cycle, Sierra and Yeager (2004) found that estimates of bank losses were very noisy, but risk rankings based on these losses were quite accurate. Our criterion for assessing jumbo CD data is analogous.

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[16] We estimated failure probabilities with the Risk-Rank model—one of two econometric surveillance models used by the Federal Reserve. See footnote 18 for more discussion of this model.

[17] We controlled for factors suggested by academic literature, examiner interviews, and specification tests. These factors included term-to-maturity, the rate on Treasury securities with comparable maturities, economic conditions (dummies for quarters and states in the union), power in local deposit markets (dummy for banks operating in an MSA), access to parent-company support (dummy for banks in holding companies), and demand for funding in excess of local supply (dummy for banks with brokered deposits). The estimation sample included only satisfactory banks to parallel the performance tests of downgrade probability and jumbo CD rankings. Confining the analysis to 1- and 2-rated banks may seem odd, akin to testing risk sensitivities of AA or better corporate debt, but there are theoretical as well as practical justifications. Managers of nonregulated firms operate with considerable latitude up to the point of bankruptcy. Bank managers, in contrast, lose much of their discretion when an unsatisfactory rating is assigned. So market data for 3-, 4-, and 5-rated banks contain assessments of ongoing supervisory intervention as well as inherent risk. Excluding unsatisfactory banks also produces more relevant evidence about the surveillance value of market feedback. Supervisors continuously monitor these institutions, so market data are unlikely to yield new information. But knowledge of deteriorating 1s and 2s would be valued because these banks do not face constant scrutiny between exams.

Surveillance value is measured not by the precision of estimated sensitivities to bank risk, but rather by the improvement in risk rankings traceable to jumbo CD yields and runoff.[18]

3. MARKET ASSESSMENTS OF RISK: THE JUMBO CD RANKINGS

The first step in assessing the value of jumbo CD data was obtaining default premiums for all sample banks with satisfactory supervisory ratings. We created two measures—a "simple" and a "complex" default premium—to reduce the likelihood that performance tests would be biased by one, possibly poor, proxy. At the root of each measure was average yield—the ratio of jumbo CD interest expense to average balance, computed with Call Report data for each bank in each quarter. To convert yields into simple default premiums, we adjusted for the average maturity of a bank's jumbo CD portfolio. To obtain a complex premium series, we used regression analysis to adjust yields for maturity and nonmaturity factors likely to affect jumbo CD rates.[19] Simple

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[18] Besides measurement error, there are several idiosyncratic aspects of the jumbo CD market that might weaken risk pricing. Jumbo CD holders often receive other bank services—loan commitments and checking accounts, for example—so the issuer might price the relationship comprehensively. Another potential explanation is that many jumbo CDs are held by state or local governments and are, therefore, practically risk-free. (Most states require banks to "pledge" Treasury or agency securities against uninsured public deposits, thereby eliminating all but fraud risk.) Still another possibility is that many banks no longer fund at the margin with jumbo CDs—these instruments are now essentially core deposits because of the declining cost of commercial paper issuance and the increasing availability of Federal Home Loan Bank advances. A final, related possibility is that posted jumbo CD rates are sticky, "clustering" around integers and even fractions like retail CD rates (Kahn, Pennacchi, and Sopranzetti 1999). These market characteristics may account for modest risk sensitivities in the yield and runoff regressions. Still, evidence presented in this section suggests the data contain information about bank condition, thereby satisfying the necessary condition for jumbo CD risk rankings to add value in surveillance.

[19] Default premiums were obtained with maturity and nonmaturity controls from the Call Report. The reporting convention for maturities changed in the middle of our sample. From 1989 to 1997, the FFIEC required banks to slot jumbo CDs in one of four buckets: "less than 3 months remaining," "3 months to 1 year remaining," "1 to 5 years remaining," and "over 5 years remaining." In 1997, the two longest maturity buckets became "1 to 3 years remaining" and "over 3 years remaining." These maturity measures are crude—jumbo CDs in the shortest bucket might have been issued years ago—but they offer the only means of controlling for term structure. We produced simple premiums by first multiplying each bank's jumbo CD balance for each maturity bucket by that quarter's yield on Treasuries of comparable maturity. The sum of the resulting values, divided by average jumbo CD balances, approximated that bank's risk-free yield. Simple default premiums for a quarter were then the difference between a bank's risk-free yield and its average jumbo CD yield that quarter. Complex premiums controlled for other factors likely to affect jumbo CD demand or supply. Specifically, average yields were regressed on average jumbo CD maturity, maturity-weighted Treasury yield (the portion of a sample bank's CDs in each maturity bucket, multiplied by that quarter's yield on a comparable-maturity Treasury), and the same nonmaturity controls used in the data-check equations in Section 3. Regression residuals served as the complex premium series. Carefully controlling in this way for maturity and nonmaturity influences on yields should render the resulting default premium series a cleaner measure of default risk.

and complex default premiums were highly correlated, exhibiting an average year-by-year correlation coefficient of 0.88.

The second step was generating a deposit-runoff series for all banks with satisfactory ratings. When significant transaction or information frictions are present, jumbo CD holders are apt to withdraw funds as failure probability rises (Park and Peristiani 1998). Another reason to examine runoff is that a bank's demand for jumbo CDs could depend on its condition. Billett, Garfinkel, and O'Neal (1998) and Jordan (2000) have documented a tendency for risky banking organizations to substitute insured for uninsured deposits to escape market discipline. If such substitution is important, escalating risk would show up in declining jumbo CD balances rather than rising default premiums. To explore these possibilities, we again computed two measures of runoff: "simple" and "complex." Simple deposit runoff was defined for each sample bank as the quarterly percentage change in jumbo CD balances.[20] The complex series was constructed by adjusting simple runoff with the same approach used to identify complex default premiums—that is, regressions of quarterly deposit runoff on maturity and nonmaturity factors likely to affect jumbo CD demand or supply. The correlation coefficient for simple and complex runoff was 35 percent, somewhat less than the correlation between simple and complex default premiums.

4. THE SURVEILLANCE BENCHMARK—DOWNGRADE PROBABILITY RANKINGS

Since the 1980s, econometric models have played an important role in bank surveillance at all three federal supervisory agencies.[21] We benchmark the performance of these models with the CAMELS-downgrade model developed by Gilbert, Meyer, and Vaughan (2002).[22] This model is a probit regres-

[20] Technically, a positive number implies growth while a negative number implies runoff. To simplify, we refer to all percentage changes as runoff. By our nomenclature, a bank can experience positive or negative jumbo CD runoff.

[21] Since the early 1990s, the Federal Reserve has relied on two econometric models, collectively known as SEER—the System for Estimating Examination Ratings. One model, the Risk-Rank model, exploits quarterly Call Report data to estimate the probability of failure over the next two years. The other model, the Ratings model, produces "shadow" CAMELS ratings—that is, the composite that would have been assigned had an examination been performed using the latest Call Report submission. Every quarter, analysts at the Board of Governors feed the data into the SEER models and forward the results to the Reserve Banks. The surveillance unit at each Bank, in turn, follows up on flagged institutions. The FDIC and the OCC use similar approaches in off-site monitoring of the banks they supervise (Reidhill and O'Keefe 1997).

[22] The model is discussed in detail here because it is possible in-sample performance has deteriorated since the Gilbert, Meyer, Vaughan (2002) estimation sample ended in 1996. Such deterioration would bias performance tests in this research in favor of the jumbo CD rankings. So, we explain the rationale for the explanatory variables and present evidence of in-sample fit to make the case that the CAMELS-downgrade model is still a good benchmark for current surveillance practices.

sion estimating the likelihood a bank with a satisfactory supervisory rating (a CAMELS 1 or 2 composite) will migrate to an unsatisfactory rating (a 3, 4, or 5 composite) in the coming eight quarters. Explanatory variables were selected in 2000 based on a survey of prior research and interviews with safety-and-soundness examiners. Table 3 describes the independent variables, as well as the expected relationship between each variable and downgrade risk. Table 5 contains summary statistics for these variables. Most variables are financial performance ratios related to leverage risk, credit risk, and liquidity risk—three risks that have consistently produced financial distress in commercial banks (Putnam 1983; Cole and Gunther 1998).

We benchmark current surveillance procedures with a CAMELS-downgrade model. Traditionally, the most popular econometric surveillance tool has been a failure-prediction model. But failures have been rare since the early 1990s, preventing re-estimation of these models. Any resulting "staleness" in coefficients could bias performance tests by compromising the surveillance benchmark used to assess jumbo CD data. Unlike failures, migration to unsatisfactory ratings remains common, so a downgrade model can be updated quarterly. (Table 4 contains 1992–2005 data on downgrade frequency.) Recent research confirms that a CAMELS-downgrade model would have improved slightly over a failure-prediction model in the 1990s (Gilbert, Meyer, and Vaughan 2002). Even more important, a downgrade model is best suited to support current supervisory practice. Institutions with unsatisfactory ratings represent significant failure risks; supervisors watch them closely and constantly to ensure progress toward safety and soundness. Most 1- and 2-rated banks, in contrast, are monitored between exams through quarterly Call Report submissions. As noted, early supervisory intervention improves chances for arresting financial deterioration. So a tool that more accurately flags deteriorating banks with Call Report data would yield the most surveillance value. These considerations have prompted one Federal Reserve Bank to "beta test" a CAMELS-downgrade model in routine surveillance and the Board of Governors to add a downgrade model to the System surveillance framework in 2006.

The CAMELS-downgrade model relies on six measures of credit risk, the risk that borrowers will not render promised interest and principal payments. These measures include the ratio of loans 30 to 89 days past due to total assets, the ratio of loans over 89 days past due to total assets, the ratio of loans in nonaccrual status to total assets, the ratio of other real estate owned to total assets (OREO), the ratio of commercial and industrial loans to total assets, and the ratio of residential real estate loans to total assets. High past-due and nonaccruing loan ratios increase downgrade probability because, historically, large portions of these loans have been charged off. OREO consists primarily of collateral seized after loan defaults, so a high OREO ratio signals poor credit-risk management. Past due loans, nonaccruing loans, and OREO are backward looking; they register asset quality problems that have already emerged (Morgan and Stiroh 2001). The ratio of commercial and industrial loans to total assets is forward looking because, historically, losses on these loans have been relatively high. The ratio of residential real estate loans to total assets also provides a forward-looking dimension because, historically, the loss rate on mortgages has been relatively low. Other things equal, an increase in dependence on commercial loans or a decrease in dependence on mortgage loans should translate into greater downgrade risk.

The model contains two measures of leverage risk—the risk that losses will exceed capital. Measures of leverage risk include the ratio of total equity (minus goodwill) to total assets and the ratio of net income to average assets (or, return on assets). Return on assets is part of leverage risk because retained earnings are an important source of capital for many banks, and higher earnings provide a larger cushion for withstanding adverse economic shocks (Berger 1995). Increases in capital protection or earnings strength should reduce the probability of migration to an unsatisfactory rating.

Liquidity risk, the risk that loan commitments cannot be funded or withdrawal demands cannot be met at a reasonable cost, also figures in the CAMELS-downgrade model. This risk is captured by two ratios: investment securities as a percentage of total assets and jumbo CD balances as a percentage of total assets. A large stock of liquid assets, such as investment securities, indicates a strong ability to meet unexpected funding needs and, therefore, should reduce downgrade probability. Liquidity risk also depends on a bank's reliance on non-core funding, or "hot money." Non-core funding, which includes jumbo CDs, can be quite sensitive to changes in money market rates. Other things equal, greater reliance on jumbo CDs implies greater likelihood of a funding runoff or an interest expense shock and, hence, a larger risk of receiving a 3, 4, or 5 rating in a future exam.

Finally, the model uses three control variables to capture downgrade risks not strictly associated with current financials. These controls include the natural logarithm of total assets because large banks are better able to reduce risk by diversifying across product lines and geographic regions. As Demsetz and Strahan (1997) have noted, however, such diversification relaxes a constraint, enabling bankers to assume more risk, so the ex ante relationship between asset size and downgrade probability is ambiguous. We also add a dummy variable for 2-rated banks because they migrate to unsatisfactory status more often than 1-rated banks. (See Table 4 for supporting evidence.) The list of control variables rounds out with a dummy for banks with management component ratings higher (weaker) than their composite rating. In such banks, examiners have raised questions about managerial competence, even though problems have yet to appear in financial statements.

We estimated the CAMELS-downgrade model for 13 overlapping two-year windows running from 1990–1991 to 2002–2003.[23] Each equation regressed downgrade incidence (1= downgraded, 0= not downgraded) in years $t + 1$ and $t + 2$ on accounting and supervisory data for banks with satisfactory ratings in the fourth quarter of year t. For example, to produce the first equation (1990–1991 in Table 6), downgrade incidence in 1990–1991 was regressed on 1989:Q4 data for all 1- and 2-rated banks that were not de novos. We continued with this timing convention, estimating equations year by year, through a regression of downgrade incidence in 2002–2003 on 2001:Q4 data. Observations ranged from 6,367 (2002–2003 equation) to 8,682 (1995–1996 equation); the count varied because bank mergers and supervisory reassessments altered the number of satisfactory institutions over the estimation period.

The model fit the data relatively well throughout the estimation sample. (Table 6 contains the results.)[24] The hypothesis that model coefficients jointly equaled zero could be rejected at the 1 percent level for all 13 equations. The pseudo-R2, the approximate proportion of variance in downgrade/no downgrade status explained by the model, was in line with numbers in prior early warning studies—ranging from 15.0 percent (1994–1995 equation) to 22.6 percent (1991–1992 equation). Estimated coefficients for seven explanatory variables—the jumbo-CD-to-total-asset ratio, the past due and nonaccruing loan ratios, the net-income-to-total-asset ratio, and the two supervisor rating dummies—were statistically significant with expected signs in all eight equations. The coefficient on the logarithm of total assets had a mixed-sign pattern, which is not surprising given ex ante ambiguity about the relationship between size and risk. The coefficients on the other six explanatory variables were statistically significant with the expected sign in at least three equations.

Comparing out-of-sample performance of jumbo CD and downgrade probability rankings is not as biased as it may first appear. True, jumbo CD rankings draw on one variable—either default premiums or deposit runoffs—while the downgrade probability rankings draw on 13 variables. But theory suggests

[23] Gilbert, Meyer, and Vaughan (2002) estimated the model for six windows running from 1990–1991 to 1995–1996. We re-estimated the model for these windows because Call Report data have since been revised, which implies slight changes in coefficients. We also wanted to use a consistent approach and consistent data for the entire estimation sample to insure subsequent out-of-sample tests of jumbo CD data were not biased against the surveillance benchmark.

[24] This table presents the results of probit regressions of downgrade status on financial-performance ratios and control variables. The dependent variable equals "1" for a downgrade and "0" for no downgrade in calendar years $t+1$ and $t+2$. Values for independent variables are taken from the fourth quarter of year t. Standard errors appear in parentheses below the coefficients. One asterisk denotes statistical significance at the 10-percent level, two at the 5-percent level, and three at the 1-percent level. The pseudo-R2 indicates the approximate proportion of variance in downgrade status explained by the model. Overall, the downgrade-prediction model fit the data well. For all eight regressions, the hypothesis that all model coefficients equal zero could be rejected at the 1-percent level of significance. In addition, eight of the 13 regression variables are significant with the predicted sign in all eight years, and all variables were significant in at least some years.

premiums and runoff should summarize overall bank risk, not just one type of exposure such as leverage or credit risk. Put another way, jumbo CD holders should sift through all available information about the condition of the issuing bank, note any changes in expected losses, and react to heightened exposures by demanding higher yields or withdrawing funds. This process should impound all relevant information—financial as well as anecdotal—into default premiums and deposit runoff just as the econometric model impounds all relevant Call Report data into a CAMELS-downgrade probability.

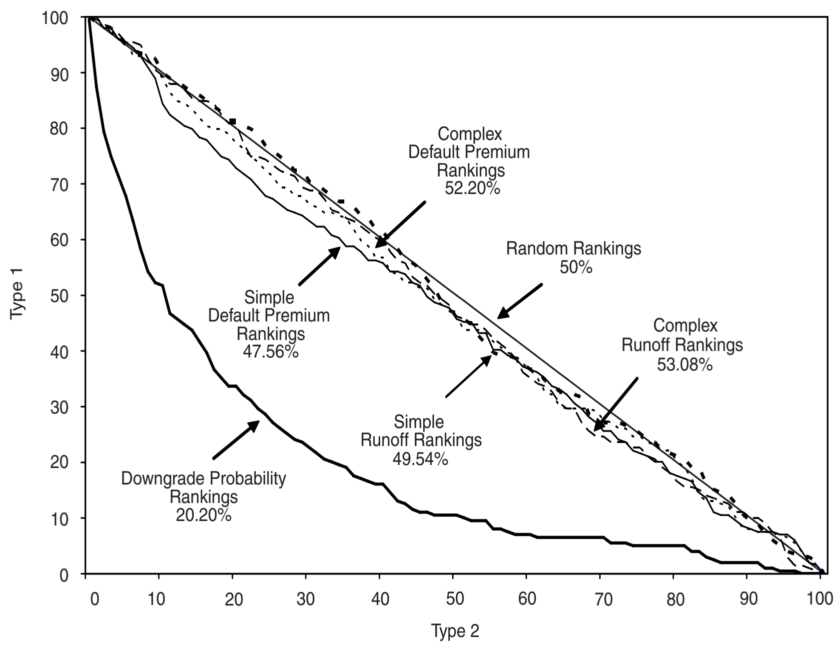
5. ASSESSING OUT-OF-SAMPLE PERFORMANCE: POWER CURVE AREAS

We assessed out-of-sample performance using both Type 1 and Type 2 error rates. Both forecast errors are costly. A missed downgrade to unsatisfactory status—Type 1 error—is costly because accurate downgrade predictions give supervisors more warning about emerging problems. A predicted downgrade that does not materialize—Type 2 error—is costly because unwarranted supervisory intervention wastes scarce examiner resources and disrupts bank operations. A tradeoff exists between the two errors—supervisors could eliminate overprediction of downgrades by assuming no banks are at risk of receiving an unsatisfactory rating in the next two years.

For each risk ranking, it is possible to draw a power curve indicating the minimum achievable Type 1 error rate for any desired Type 2 error rate (Cole, Cornyn, and Gunther 1995). For example, tracing the curve for simple default premium rankings starts by assuming no sample bank is a downgrade risk. This assumption implies all subsequent downgrades are surprises—a 100 percent Type 1 error rate. Because no banks are incorrectly classified as downgrade risks, the Type 2 error rate is zero. The next point on the curve is obtained by selecting the bank with the highest simple default premium (maturity-adjusted spread over Treasury). If that bank suffers a downgrade in the following eight quarters, then the Type 1 error rate decreases slightly. The Type 2 error rate remains zero because, again, no institutions are incorrectly classified as downgrade risks. If the selected bank does not suffer a downgrade, then the Type 1 error rate remains 100 percent, and the Type 2 error rate increases slightly. Selecting banks from highest to lowest default premium and recalculating error rates each time produces a power curve. At the lower right extreme of the curve, all banks are considered downgrade risks—the Type 1 error rate is 0 percent, and the Type 2 error rate is 100 percent. Figure 1 illustrates with the power curves for downgrade probability and jumbo CD rankings for the 1992–1993 test window.

Areas under power curves provide a basis for comparing out-of-sample performance across risk rankings. The area for each ranking is expressed as a percentage of the total area of the box. A smaller percentage implies a lower

Figure 1 How Well Do Jumbo CD and Downgrade Probability Rankings Perform Out-of-Sample? 1992–1993 Test Window



Notes: This Figure depicts the power curves for risk rankings based on jumbo CD default premiums, jumbo CD runoff, and downgrade probabilities (as produced by the CAMELS-downgrade model) for the 1992–1993 out-of-sample test window. These curves reflect the tradeoff between Type 1 and Type 2 errors. (Type 1 errors are missed downgrades, and Type 2 errors are overpredicted downgrades.) A convenient way to compare the performance of risk rankings is to calculate the area under the power curve for each ranking and express that area as a percentage of the total for the box. Smaller areas are desired because they imply a simultaneous reduction in both types of errors. The 50-percent line is the power curve produced when downgrade risks are selected randomly over a large number of trials. The power curves above show that rankings based on downgrade probabilities would have significantly outperformed rankings based on jumbo CD default premiums and runoff. Indeed, jumbo CD rankings would not have improved materially over random rankings.

overall Type 1 and Type 2 error rate and, hence, a more accurate forecast. The area for a "random-ranking" power curve offers an example as well as a yardstick for evaluating the economic significance of differences in forecast accuracy. Random selection of downgrade candidates, over a large number of trials, will produce power curves with an average slope of negative one. Put another way, the area under the random-ranking power curves, on average, equals 50 percent of the total area of the box. Power curve areas can be compared—jumbo CD ranking against downgrade probability rankings or either ranking against a random ranking—for any error rate. Assessing forecast accuracy this way, though somewhat atheoretical, makes best use of existing data. A more appealing approach would minimize a loss function explicitly weighing the benefits of early warning about financial distress against the costs of wasted examination resources and unnecessary regulatory burden. Then, the relative performance of risk rankings could be assessed for the optimal Type 1 (or Type 2) error rate. The requisite data, however, are not available.

A specific example will clarify the mechanics of the "horse race" we run for risk rankings. To assess the surveillance value of simple default premiums for 1992–1993, we start by assuming it is early 1992, just after fourth quarter 1991 data became available. In accordance with standard surveillance procedures, 1990–1991 downgrade incidences are regressed on 1989:Q4 data for the 13 explanatory variables in the CAMELS-downgrade model. Model coefficients are then applied to 1991:Q4 data to estimate the probability that each 1- and 2-rated bank will migrate to an unsatisfactory condition between 1992:Q1 and 1993:Q4. These banks are then ranked from highest to lowest downgrade probability. At the same time, all banks with satisfactory supervisory ratings are ranked from highest to lowest simple default premium (maturity adjusted spread over Treasury), also using 1991:Q4 data, under the assumption that high spreads map into high downgrade probabilities. After two years, the record of missed and overpredicted downgrades is compiled to generate power curve areas for each ranking. A smaller area for the downgrade probability ranking would imply that simple default premiums added no surveillance value in the 1992–1993 test window.

6. EMPIRICAL EVIDENCE

Downgrade Model Rankings and Jumbo CD Rankings—Full-Sample Results

The evidence suggests jumbo CD default premiums would have contributed nothing to bank surveillance between 1992 and 2005 when used to forecast downgrades two years out. (Columns 2, 3, and 4 of Table 7 contain the relevant power curve areas.) Over the 13 test windows, the average area under the simple default premium power curve (45.63 percent) and the average area under the complex default premium power curve (49.70 percent) did not differ statistically or economically from the random-ranking benchmark (50 percent). In contrast, the average area under the downgrade model power curve (19.66 percent) came to less than half of that benchmark. Power curve areas for individual two-year test windows showed the same patterns. Specifically, downgrade model areas ranged from 15.24 percent (1996–1997) to 22.39 percent (1994–1995); simple default premium areas ran from 41.56 percent (2003–2004) to 50.57 percent (1994–1995); and complex default premium areas varied from 45.69 percent (1998–1999) to 52.20 percent (1992–1993). The poor performance of jumbo CD rankings relative to downgrade model rankings suggests default premiums would not have flagged banks missed by conventional surveillance. The poor performance relative to the random-selection benchmark suggests default premiums would not have increased supervisor confidence about rankings produced by the CAMELS-downgrade model.

Out-of-sample performance of risk rankings based on jumbo CD runoff was no better. (Columns 5 and 6 of Table 7 contain the relevant power curve areas.) The average area for simple runoff rankings across all test windows was 46.12 percent while the average for complex runoff was 50.47 percent—again, statistically and economically indistinguishable from random selection. And once again, patterns were consistent across individual two-year test windows. Power curve areas for simple runoff rankings varied from 43.56 percent (1999–2000) to 49.54 percent (1992–1993), areas under complex runoff curves from 45.75 percent (1998–1999) to 53.08 percent (1992–1993). This consistently poor performance suggests runoff rankings would not have helped spot downgrade risks two years out between 1992 and 2005.

Changing forecast horizons did not alter the results. Over 13 one-year windows, downgrade model rankings produced an average power curve area of 17.37 percent (standard deviation across test windows of 2.39 percent). In contrast, simple default premium rankings produced an average area of 45.38 percent (standard deviation of 2.95 percent) and complex default premium rankings, an average area of 50.29 percent (standard deviation of 3.26 percent). Areas for runoff rankings were even closer to the random-ranking benchmark—46.79 percent on average for simple runoff (standard deviation across the 13 windows of 2.47 percent) and 50.87 percent for complex runoff (standard deviation of 3.45 percent). Lengthening the forecast horizon to three years yielded similar numbers. This evidence goes to the timeliness of information in jumbo CD rankings. As noted, market data could enhance surveillance by flagging problems before existing tools. But, between 1992 and 2005, jumbo CD data would not have improved over random selection at any forecasting horizon, much less current surveillance procedures. Put another way, feedback from the jumbo CD market would not have provided supervisors with earlier warning about developing problems.

Jumbo CD rankings constructed from both default premiums and runoff did not improve over random selection, either. In theory, price and quantity signals from the jumbo CD market, though weak when used singly, could jointly capture useful information about future bank condition. If so, a model relying only on multiple signals could add value—even if performance relative to the benchmark was poor—by reducing supervisor uncertainty about banks flagged by conventional surveillance. We explored this possibility by estimating a downgrade model with (1) only simple default premiums and runoff as explanatory variables and (2) only complex default premiums and runoff as explanatory variables. Out-of-sample performance was then tested over a variety of forecasting horizons for 1992–2005. Columns 7 and 8 in Table 7 contain the results for two-year windows; they are representative. Over all 13 tests, the power curve area for the bivariate "simple" model averaged 45.55 percent (standard deviation across individual windows of 2.49 percent) and 50.05 percent for the bivariate "complex" model (standard deviation of 2.99 percent). Further perspective can be gained by comparing these numbers to power curve areas produced by a "pared down" model including only the dummy variables in the baseline CAMELS-downgrade model. The average power curve area for this model across the 13 two-year windows was 30.07 percent. Taken together, this evidence suggests jumbo CD data would not have reduced supervisory uncertainty about banks flagged by conventional surveillance tools.

Downgrade Model Rankings and Jumbo CD Rankings—Sub-Sample Results

Although jumbo CD risk rankings would not have contributed to general surveillance of 1- and 2-rated banks, default premiums and runoff might improve monitoring of specific cohorts such as banks with short jumbo CD portfolios, large asset portfolios, no foreign deposits, low capital ratios, or significant "deposits at risk."

The marginal-average problem noted earlier could in part account for the weak performance of default premium rankings. As an arithmetic matter, today's average yield will be more representative of today's risk levels if jumbo CD maturities are short. To explore this possibility, we replicated all out-of-sample tests described earlier in Section 6 on a sub-sample of banks with weighted-average portfolio maturities under six months. The results did not change. At the two-year horizon, for example, the average area under complex default premium power curves was 42.97 percent. At the one-year horizon, the average area was 42.30 percent. Put simply, long jumbo CD maturities do not account for the poor performance of default premiums.

The jumbo CD market might emit stronger risk signals for large, complex banking organizations. Jumbo CDs at community banks may be more like core deposits than money market instruments. And because prices and quantities of core deposits are known to be sticky (Flannery 1982), yields and runoff of community bank jumbos could respond sluggishly to changes in risk no matter how short the maturity of the portfolio. Another reason large bank signals may be more informative is that monitoring costs for their uninsured depositors are lower—these institutions have publicly traded securities and are closely followed by market analysts. To test for an asset-threshold effect, we reproduced all out-of-sample tests from Section 6 on a sub-sample of banks holding more than the median level of assets. Out-of-sample results for this sub-sample were qualitatively similar to the results from the full sample. Across the 13 two-year test windows, for example, the area under the simple default premium power curves averaged 44.90 percent (compared with 45.63 percent for the full sample). We also tested risk rankings for banks holding more than $1 billion in assets and for banks with SEC registrations. Each time, we compared results from the large bank sub-sample with results from the remaining sub-sample (i.e., banks holding less than median assets, banks holding less than $1 billion in assets, and banks with no SEC registration), looking for performance differences across size cohorts. Size-split evidence was consistent: for large as well as community banks, the CAMELS-downgrade model proved to be the far superior surveillance tool, and rankings based on default premiums and runoff barely improved over random rankings.

Jumbo CD default premiums and runoff might improve off-site monitoring of banks with no foreign deposits. The National Depositor Preference Act of 1993 elevated claims of domestic depositors over claims of foreign depositors, reducing expected losses for jumbo CD holders (Marino and Bennett 1999). Domestic holders of jumbo CDs issued by banks with foreign offices may have perceived no default-risk exposure because of the financial cushion provided by foreign deposits. To test for a depositor-preference effect, we screened out banks with foreign deposits and replicated all out-of-sample tests. Again, the results mirrored the full-sample results; for example, for the two-year test windows, the average power curve area under the simple default premium rankings was 45.70 percent, virtually unchanged from the full sample (45.63 percent). Even for banks with no foreign-deposit cushion, jumbo CD rankings contained no useful supervisory information.

Finally, the jumbo CD market might yield clues about emerging problems in banks with high levels of uninsured deposits or low levels of capital. In theory, jumbo CD holders with more exposure—either because their uninsured balances are high or bank capital levels are low—have greater incentive to monitor and discipline risk. So we produced rankings for the quartile of sample banks with the largest volume of "deposits at risk" and the quartile with the lowest ratios of equity-to-assets (adjusted for bank size). Again, default premium and runoff rankings did not improve over random selection, much less conventional surveillance. As a final check, we looked at various intersections of the sub-samples—banks with high deposits at risk *and* low capital, banks with no foreign deposits *and* short jumbo CD maturities, etc. We generated rankings based on default premiums, deposit runoff, and both default premiums and deposit runoff. The results across all tests were consistent—jumbo CD rankings did not improve materially over random rankings at any forecast horizon.

Default Premiums and Runoff as Regressors in the Downgrade Model

Although default premium and runoff perform poorly as independent risk signals, they could add value as regressors in the CAMELS-downgrade model. Indeed, previous research has identified surveillance ratios with this property (Gilbert, Meyer, and Vaughan 1999). To pursue this angle, we estimated an "enhanced" CAMELS-downgrade model, adding both simple and complex measures of premiums and runoff to the 13 baseline explanatory variables. As before, out-of-sample performance was gauged by impact on power curve areas—first when default premiums and runoff were added to the baseline model and then when these variables were dropped from the enhanced model. As a further check, we assessed performance with the quadratic probability score (QPS)—a probit analogue for root mean square error (Estrella 1998; Estrella and Mishkin 1998).[25] If default premiums and runoff enhance the CAMELS-downgrade model, removing them from the enhanced model will boost QPS. Columns 9 and 10 in Table 7 contain power curve areas for the simple and complex enhancements of the downgrade model. Column 2 of panel A in Table 8 notes the impact of the two simple jumbo CD series on power curve areas; column 2 of panel B in Table 8 shows the impact of the series on QPS. (Results for complex default premiums and runoff are not reported because they mirror results for the simple series.) To facilitate interpretation, we note the impact on QPS and power curve areas of other variable blocks—such as the leverage-risk variables (equity-to-asset ratio and return on assets) and control variables (log of total assets, dummy for composite rating of 2, and dummy for management component rating weaker than the composite rating)—in columns 3 through 6 of panels A and B in Table 8. In Table 6, changes in QPS and power curve areas are expressed in percentage-change terms to permit direct comparison.

In performance tests for 1992–2005, default premiums and runoff did not enhance the CAMELS-downgrade model. Adding simple versions of the series increased (worsened) average power curve area by 4.17 percent (0.82 percentage points, from 19.66 percent for the baseline model to 20.48 percent). Removing these series from the enhanced downgrade model improved performance slightly by the power curve metric (reduced average power curve area by 0.26 percent) and worsened performance even more slightly by the QPS metric (increased average QPS by 0.06 percent). The leverage-risk variables provide

[25] To obtain QPS, we first computed downgrade probability for each sample bank with the CAMELS-downgrade model. Then, we subtracted R_t—a binary variable equal to one if the bank was downgraded in the out-of-sample window and zero if not—from the downgrade probability estimate. Finally, we squared the difference, multiplied the result by two, and averaged across all sample banks. An ideal model generates probabilities close to unity for banks with subsequent downgrades and probabilities close to zero for non-downgrades, so higher QPS figures imply weaker out-of-sample performance, just as higher power curve areas imply weaker performance.

some perspective on the economic significance of these changes—dropping both of them increased the power curve area (worsened performance) by an average of 5.20 percent and increased the average QPS (worsened performance) by an average of 2.03 percent. These results held up in tests on the various sample cuts and forecasting horizons described in the previous subsections.

7. DISCUSSION

It is possible that some combination of measurement error and idiosyncrasies in the jumbo CD market accounts for our results. These factors may not be important enough to remove all evidence of risk pricing from jumbo CD data, but they may be important enough to prevent risk rankings based on the data from imparting valuable surveillance information.

But data problems and market frictions are unlikely to explain away the findings. As noted, recent studies using actual debt and equity market data rather than accounting proxies have found only modest surveillance value in market signals. Rather, the economic environment since the early 1990s probably plays an important role. Over this period, bank profitability and capital ratios soared to record highs. Some economists attribute these trends to an unprecedented economic boom that allowed banks to reap the upside of expansions into risky new markets and product lines (Berger et al. 2000). Others argue that stakeholders of large complex banking organizations insisted on greater capital cushions because of increasingly sophisticated risk exposures (Flannery and Rangan 2002). In such a high-profit, high-capital environment, jumbo CD signals—no matter how accurately measured or precisely determined—would convey little information because the benefits of monitoring are so low. Such an explanation would account for the successful use of average yields in bank-risk studies on data from the 1980s—a time when financial distress was fairly common and failures were sharply rising. Such an explanation would also account for the evidence in Martinez-Peria and Schmukler (2001). With a data set and research strategy similar to ours, they studied the impact of banking crises on market discipline in Argentina, Chile, and Mexico, finding little discipline before, but significant discipline after, the crises.

8. CONCLUSION

The evidence suggests that feedback from the jumbo CD market would have added no value in bank surveillance between 1992 and 2005. Throughout the decade, risk rankings produced by a CAMELS downgrade—a model chosen to benchmark current surveillance practices—would have significantly outperformed risk rankings based on default premiums and runoff. Moreover, jumbo CD rankings would have improved little over random orderings.

Finally, adding jumbo CD signals to the downgrade model would not have improved its out-of-sample performance. These results hold up for a variety of sample cuts and forecast horizons. Taken together, these results imply that the marginal surveillance value of jumbo CD signals is less than the marginal production cost—even if that cost is very low.

Our results carry mixed implications for proposals to incorporate market data more formally into bank supervision. On the one hand, the evidence suggests available jumbo CD data would do little to enhance surveillance, thereby clearing the way for experimentation with other, "purer" market signals. On the other hand, if the "unique sample period" explanation for our results is true, then it is likely the surveillance value of signals from the market for bank debt and equity will vary over time. Other things equal, such time variation would lower the net benefit of integrating market data into current surveillance routines. Interpreted in this light, our findings imply that future policy and research work on market data should focus on identifying the specific bank claims that yield the most surveillance value in each state of the business cycle. Put another way, our findings—when viewed with other recent research—suggest the supervisory return from reliance on a single market signal through all states of the world may have been overestimated.

Table 1 Do Prior Studies Point to Risk Pricing in the Jumbo CD Market?

Authors	Issuer of Jumbo CD	Country	Sample Dates	Yield or Runoff?	Risk Pricing?
Crane (1979)	Bank	United States	1974	Yield	Somewhat
Goldberg & Lloyd-Davies (1985)	Bank	United States	1976–1982	Yield	Yes
Baer & Brewer (1986)	Bank	United States	1979–1982	Yield	Yes
Hannan & Hanweck (1988)	Bank	United States	1985	Yield	Yes
James (1988)	Bank	United States	1984–1986	Yield	Yes
Cargill (1989)	Bank	United States	1984–1986	Yield	Yes
James (1990)	Bank	United States	1986–1987	Yield	Yes
Keeley (1990)	Bank	United States	1984–1986	Yield	Yes
Ellis & Flannery (1992)	Bank	United States	1982–1988	Yield	Yes
Cook & Spellman (1994)	Thrift	United States	1987–1988	Yield	Yes
Crabbe & Post (1994)	Bank	United States	1986–1991	Runoff	No
Brewer & Mondschean (1994)	Thrift	United States	1987–1989	Yield	Yes
Park (1995)	Bank	United States	1985–1992	Both	Yes
Park & Peristiani (1998)	Thrift	United States	1987–1991	Both	Yes
Jordan (2000)	Bank	United States	1989–1995	Both	Yes
Martinez Peria & Schmuckler (2001)	Bank	Argentina, Chile, Mexico	1981–1997	Both	Yes
Goldberg & Hudgins (2002)	Thrift	United States	1984–1994	Runoff	Yes
Birchler & Maechler (2002)	Bank	Switzerland	1987–2001	Runoff	Yes
Maechler & McDill (2003)	Bank	United States	1987–2000	Runoff	Yes
Hall, King, Meyer, & Vaughan (2005)	Bank	United States	1988–90, 1993–95	Both	Yes

Notes: This table summarizes the literature on risk pricing by jumbo CD holders. ("Bank" refers to commercial banks, bank holding companies, and thrift institutions. "Risk pricing" refers to price or quantity responses to a change in bank condition.) These studies used both cross-section and time-series techniques along with a variety of risk proxies and control variables. The weight of the evidence indicates bank condition is priced, suggesting that jumbo CD data might add value in off-site surveillance.

Table 2 Do Jumbo CD Data Contain Evidence of Risk Pricing? Evidence from Regressions of Yields and Runoff on Failure Probabilities

Sensitivity of Jumbo CD Yields and Runoff to Failure Probability 1988: Q1–2004:Q4

Independent Variable	Dependent Variable: Yields Coefficient (Std. Error)	Dependent Variable: Runoff Coefficient (Std. Error)
Failure Probability	0.0108***	-0.3309***
(Lagged One Quarter)	(0.0038)	(0.0505)
Maturity-Weighted Treasury Yield	0.6878***	0.5324
	(0.0391)	(0.5211)
Average Portfolio Maturity	0.0820***	-1.8743***
	(0.0209)	(0.2785)
Maturity-Treasury Interactive	-0.0140***	-0.3408*
	(0.0147)	(0.1959)
Holding Company Dummy	-0.0595	-0.9042***
	(0.0136)	(0.1808)
Brokered Deposit Dummy	0.1403***	-0.0868
	(0.0270)	(0.3592)
MSA Dummy	0.1384***	1.7813
	(0.0913)	(1.2164)
R^2	0.2288	0.0125
F-statistic Control Variables	394.16***	33.25***
F-statistic Time Dummies	155.69***	21.88***
F-statistic State Dummies	21.53***	15.21***
Observations	229,486	229,486

Notes: This table reports results for regressions of jumbo CD yields and runoff on failure probabilities and various controls. Equations were estimated on a sub-sample of banks with satisfactory CAMELS ratings. Asterisks denote statistical significance at the 10- (*), 5- (**), and 1- (***) percent levels. A positive and significant failure-probability coefficient in the yield equation and/or a negative and significant coefficient in the runoff equation constitute evidence that bank condition is priced. The results indicate that greater risk of failure translates, on average, into higher yields and larger runoff, suggesting that jumbo CD data may have surveillance value in our sample.

Table 3 Which Variables Predict Migration to an Unsatisfactory CAMELS Rating?

	Independent Variables	Symbol	Impact on Downgrade Risk
Credit Risk	Loans past due 30-89 days (% of total assets)	PAST DUE 30	+
	Loans past due over 89 days (% of total assets)	PAST DUE 90	+
	Loans in nonaccrual status (% of total assets)	NONACCRUING	+
	"Other real estate owned" (% of total assets)	OREO	+
	Commercial & industrial loans (% of total assets)	COMMERICAL	+
	Residential real estate loans (% of total assets)	RESIDENTIAL	–
Leverage Risk	Equity capital minus goodwill (% of total assets)	NET WORTH	–
	Net income (% of average assets)	ROA	–
Liquidity Risk	Book value of investments (% of total assets)	SECURITIES	–
	Time deposits over $100,000 (% of total assets)	JUMBO CDs	+
Controls	Natural logarithm of total assets (thousands of dollars)	SIZE	Ambiguous
	Dummy for banks with composite CAMELS rating=2	CAMELS-2	+
	Dummy for banks with management rating > composite rating	MANAGEMENT	+

Notes: This table lists the independent variables in the CAMELS-downgrade model. Signs note the hypothesized relationship between each variable and the likelihood of downgrade from a satisfactory (CAMELS 1 or 2 composite) to an unsatisfactory rating (CAMELS 3, 4, or 5 composite). For example, the negative sign on the NET WORTH variable indicates that, other things equal, higher capital levels reduce the likelihood of migration to an unsatisfactory rating over the next two years.

Table 4 How Common Is Migration to Unsatisfactory CAMELS Ratings? Evidence from 1992–2005

Year of Migration to 3, 4, or 5 Rating	Rating at Beginning of Year	Banks with 1 & 2 Rating	Number Migrating to 3, 4, or 5 Rating	Percentage Migrating to 3, 4, or 5 Rating	Total Downgrades to 3, 4, or 5 Rating
1992	1	1,959	22	1.12	425
	2	5,275	403	7.64	
1993	1	2,289	7	0.31	182
	2	5,978	175	2.93	
1994	1	2,919	9	0.31	162
	2	5,742	153	2.66	
1995	1	3,106	8	0.26	102
	2	4,905	94	1.92	
1996	1	3,295	10	0.30	127
	2	4,518	117	2.59	
1997	1	3,250	7	0.22	125
	2	3,744	118	3.15	
1998	1	3,027	19	0.63	154
	2	3,101	135	4.35	
1999	1	3,064	19	0.62	198
	2	3,041	179	5.89	
2000	1	2,843	12	0.42	195
	2	3,084	183	5.93	
2001	1	2,661	12	0.45	231
	2	3,153	219	6.95	
2002	1	2,449	11	0.45	230
	2	3,216	219	6.81	
2003	1	2,283	16	0.70	193
	2	3,101	177	5.71	
2004	1	2,111	10	0.47	124
	2	2,950	114	3.86	
2005	1	2,573	10	0.39	95
	2	3,650	85	2.33	

Notes: This table demonstrates that banks with satisfactory composites ratings (CAMELS 1 or 2) frequently migrate to unsatisfactory ratings (3, 4, or 5), thereby permitting yearly re-estimation of the CAMELS-downgrade model. The data also show that 2-rated banks are much more likely to migrate to unsatisfactory ratings than 1-rated banks.

Table 5 Selected Summary Statistics—Jumbo CD Data and Regressors for the CAMELS-Downgrade Model

	Variable	Median	Mean	Standard Deviation
Credit Risk	PAST DUE 30	0.68	0.90	0.84
	PAST DUE 90	0.07	0.21	0.39
	NONACCRUING	0.19	0.38	0.56
	OREO	0.03	0.20	0.46
	COMMERICAL	7.65	9.28	6.97
	RESIDENTIAL	14.26	15.91	10.68
Leverage Risk	NET WORTH	8.84	9.79	4.74
	ROA	1.17	1.23	2.05
Liquidity Risk	SECURITIES	28.65	30.41	14.87
	JUMBO CDs	8.00	9.33	6.56
Controls	SIZE	11.07	11.21	1.29
	CAMELS-2	1.00	0.61	0.49
	MANAGEMENT	0.00	0.18	0.39
	"Simple" Default Premium	0.47	0.42	2.83
	"Complex" Default Premium	NA	NA	2.18
	"Simple" Deposit Runoff	9.39	19.25	52.02
	"Complex" Default Premium	NA	NA	33.03

Notes: This table contains summary statistics for the independent variables used in the CAMELS-downgrade prediction model, computed over all year-end regression observations from 1989 to 2001. Summary statistics for the default premiums and deposit runoff series used in jumbo CD risk rankings are also provided for comparison. The "complex" measures of premium and runoff are regression residuals, so means and medians are not meaningful, but standard deviations are roughly in line with their "simple" counterparts. The correlation coefficients between the "simple" and "complex" measures are 88 percent for default premiums and 35 percent for the runoff.

Table 6 How Well Does the CAMELS-Downgrade Model Fit the Data? Downgrade Years: 1990–2005

	Independent Variable	**Period of Downgrade in CAMELS Rating**			
		1990–1991	1991–1992	1992–1993	1993–1994
	Intercept	-2.087***	-0.957***	-0.081	0.048
		(0.246)	(0.264)	(0.318)	(0.375)
Credit Risk	PAST DUE 30	0.112**	0.150***	0.136***	0.174***
		(0.021)	(0.022)	(0.026)	(0.033)
	PAST DUE 90	0.376***	0.328***	0.239***	0.304***
		(0.039)	(0.040)	(0.047)	(0.060)
	NONACCRUING	0.235***	0.199***	0.291***	0.178***
		(0.029)	(0.030)	(0.036)	(0.045)
	OREO	0.220***	0.216***	0.145***	0.167***
		(0.030)	(0.032)	(0.031)	(0.043)
	COMMERCIAL	0.009***	0.013***	0.009**	0.002
		(0.003)	(0.003)	(0.004)	(0.005)
	RESIDENTIAL	-0.005***	-0.004	-0.004	-0.005
		(0.002)	(0.002)	(0.003)	(0.003)
Leverage Risk	NET WORTH	-0.054***	-0.048***	-0.073***	-0.074***
		(0.010)	(0.011)	(0.013)	(0.013)
	ROA	-0.241***	-0.318***	-0.200***	-0.263***
		(0.035)	(0.039)	(0.043)	(0.051)
Liquidity Risk	SECURITIES	-0.016***	-0.017***	-0.013***	-0.009***
		(0.002)	(0.002)	(0.002)	(0.003)
	JUMBO CDs	0.017***	0.019***	0.015***	0.017***
		(0.003)	(0.003)	(0.004)	(0.005)
Controls	SIZE	0.079	-0.029	-0.125***	-0.147***
		(0.017)	(0.019)	(0.024)	(0.030)
	CAMELS-2	0.633***	0.517***	0.509***	0.432***
		(0.062)	(0.068)	(0.087)	(0.102)
	MANAGEMENT	0.488***	0.401***	0.478***	0.466***
		(0.051)	(0.054)	(0.061)	(0.069)
	Number of Observations	8,494	8,065	7,837	8,060
	Pseudo-R^2	0.219	0.226	0.209	0.161

Table 6 (Continued) How Well Does the CAMELS-Downgrade Model Fit the Data?

	Independent Variable	**Period of Downgrade in CAMELS Rating**				
		1994–1995	1995–1996	1996–1997	1997–1998	1998–1999
	Intercept	-0.780*	-0.011	-0.162	-1.371***	-1.603***
		(0.402)	(0.436)	(0.415)	(0.388)	(0.352)
Credit Risk	PAST DUE 30	0.119***	0.164***	0.093***	0.189***	0.186***
		(0.035)	(0.035)	(0.029)	(0.033)	(0.030)
	PAST DUE 90	0.296***	0.322***	0.347***	0.399***	0.182***
		(0.064)	(0.074)	(0.057)	(0.064)	(0.058)
	NONACCRUING	0.192***	0.145***	0.187***	0.157***	0.163***
		(0.046)	(0.051)	(0.044)	(0.046)	(0.044)
	OREO	0.192***	0.153***	0.156**	0.091	0.118
		(0.044)	(0.052)	(0.067)	(0.059)	(0.087)
	COMMERICAL	0.007	0.013***	0.005	0.010**	0.015
		(0.005)	(0.005)	(0.005)	(0.005)	(0.004)
	RESIDENTIAL	-0.002	-0.013***	0.000	-0.009***	-0.004
		(0.004)	(0.004)	(0.003)	(0.003)	(0.003)
Leverage Risk	NET WORTH	-0.032**	-0.034***	-0.020*	-0.036***	-0.011
		(0.014)	(0.013)	(0.012)	(0.014)	(0.010)
	ROA	-0.229***	-0.164***	-0.110**	-0.393***	-0.133
		(0.052)	(0.038)	(0.044)	(0.063)	(0.040)
Liquidity Risk	SECURITIES	-0.002	-0.010***	-0.011***	-0.015***	-0.007**
		(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
	JUMBO CDs	0.024***	0.020***	0.019***	0.023***	0.008*
		(0.005)	(0.005)	(0.004)	(0.005)	(0.004)
Controls	SIZE	-0.150***	-0.202***	-0.101***	-0.150***	-0.071***
		(0.033)	(0.035)	(0.030)	(0.032)	(0.027)
	CAMELS-2	0.594***	0.589***	0.760***	0.501***	0.716***
		(0.104)	(0.103)	(0.093)	(0.099)	(0.078)
	MANAGEMENT	0.389***	0.510***	0.535***	0.406***	0.518***
		(0.075)	(0.078)	(0.081)	(0.083)	(0.077)
	Number of Observations	8,665	8,682	8,585	8,314	7,818
	Pseudo-R^2	0.150	0.188	0.223	0.184	0.166

Table 6 (Continued) How Well Does the CAMELS-Downgrade Model Fit the Data?

	Independent Variable	Period of Downgrade in CAMELS Rating			
		1999–2000	2000–2001	2001–2002	2002–2003
	Intercept	-1 .118***	-1 .061***	-1 .788***	-1 .528***
		(0 .360)	(0 .358)	(0 .342)	(0 .342)
Credit Risk	PAST DUE 30	0 .169***	0 .184***	0 .170***	0 .171***
		(0 .029)	(0 .032)	(0 .026)	(0 .027)
	PAST DUE 90	0 .217***	0 .417***	0 .321***	0 .147**
		(0 .055)	(0 .059)	(0 .059)	(0 .061)
	NONACCRUING	0 .227***	0 .165***	0 .250***	0 .285***
		(0 .044)	(0 .045)	(0 .042)	(0 .041)
	OREO	0 .117	0 .157*	0 .175*	0 .082
		(0 .076)	(0 .087)	(0 .090)	(0 .073)
	COMMERICAL	0 .013***	0 .013***	0 .010***	0 .004
		(0 .004)	(0 .004)	(0 .004)	(0 .004)
	RESIDENTIAL	-0 .002	0 .002	0 .001	-0 .005
		(0 .003)	(0 .003)	(0 .003)	(0 .003)
Leverage Risk	NET WORTH	-0 .044***	-0 .036***	-0 .008	-0 .025**
		(0 .011)	(0 .010)	(0 .009)	(0 .010)
	ROA	-0 .199***	-0 .254***	-0 .135***	-0 .134***
		(0 .046)	(0 .043)	(0 .039)	(0 .037)
Liquidity Risk	SECURITIES	-0 .002	-0 .005**	-0 .010***	-0 .005*
		(0 .003)	(0 .003)	(0 .003)	(0 .002)
	JUMBO CDs	0 .015***	0 .017***	0 .023***	0 .015***
		(0 .004)	(0 .004)	(0 .004)	(0 .004)
Controls	SIZE	-0 .099***	-0 .106***	-0 .069***	-0 .065**
		(0 .029)	(0 .028)	(0 .026)	(0 .026)
	CAMELS-2	0 .780***	0 .799***	0 .878***	0 .797***
		(0 .079)	(0 .081)	(0 .083)	(0 .083)
	MANAGEMENT	0 .564***	0 .538***	0 .338***	0 .491***
		(0 .080)	(0 .084)	(0 .092)	(0 .088)
	Number of Observations	7 ,341	6 ,968	6 ,582	6 ,367
	Pseudo-R^2	0 .190	0 .206	0 .210	0 .184

Note: See footnote 24.

Table 7 Do Jumbo CD Default Premiums or Runoff Add Value in Bank Surveillance? Full-Sample, Two-Year Horizon

Out-of-Sample Test Window (1)	CAMELS-Downgrade Model (2)	Simple Default Premiums (3)	Complex Default Premiums (4)	Simple Runoff (5)	Complex Runoff (6)	Simple Premium + Runoff Model (7)	Complex Premium + Runoff Model (8)	Downgrade Model + Simple Premiums/ Runoff (9)	Downgrade Model + Complex Premiums/ Runoff (10)
1992–1993	20.20	47.56	52.20	49.54	53.08	50.51	52.88	21.12	21.67
1993–1994	21.81	47.10	47.48	47.22	48.62	48.31	49.30	23.22	21.07
1994–1995	22.39	50.57	49.57	45.06	50.90	46.18	49.55	22.67	19.94
1995–1996	17.51	43.50	46.81	47.42	48.49	45.52	46.48	18.64	18.87
1996–1997	15.24	46.93	49.01	46.31	51.03	46.39	48.79	18.38	16.16
1997–1998	19.24	46.34	48.12	45.01	49.33	45.22	48.92	19.72	20.59
1998–1999	21.39	47.27	45.69	47.20	45.75	47.24	45.70	21.87	21.45
1999–2000	19.55	45.62	48.83	43.56	48.94	43.41	48.64	19.33	19.19
2000–2001	18.77	44.31	51.78	46.10	52.03	45.38	52.13	20.11	18.47
2001–2002	18.92	43.82	51.78	44.93	51.85	44.40	51.90	19.64	20.23
2002–2003	19.46	45.35	51.61	46.27	51.56	44.34	51.45	19.56	19.36
2003–2004	20.36	41.56	51.30	46.12	52.54	43.02	52.30	21.18	20.51
2004–2005	20.76	43.29	51.89	44.85	52.00	42.18	52.59	20.77	21.14
All Years	**19.66**	**45.63**	**49.70**	**46.12**	**50.47**	**45.55**	**50.05**	**20.48**	**19.95**

Notes: This table summarizes evidence about the surveillance value of jumbo CD data. Each cell in columns 2 through 10 contains the area under the power curve for a specific risk-ranking produced by a specific surveillance tool over a specific test window. Smaller areas imply lower Type 1 and Type 2 error rates and, thus, better performance. Column 2 contains areas for downgrade probability rankings to benchmark current practices. Columns 3 through 10 contain rankings based on various uses of jumbo CD default premiums and runoff. The evidence suggests the data would have added no value in surveillance between 1992 and 2005. Risk rankings produced by the CAMELS-downgrade model (column 2) performed considerably better than random rankings (average power curve area of 50 percent). But, rankings based on default premiums or runoff (columns 3 through 6) as well as rankings based on both series (columns 7 and 8) barely outperformed random rankings. Finally, default premiums and runoff (columns 9 and 10) did not improve out-of-sample performance of the CAMELS-downgrade model.

Table 8 Do Jumbo CD Default Premiums or Runoff Enrich the CAMELS-Downgrade Model?

Panel A: Percentage Change in Power Curve Area

Out-of-Sample Window (1)	Default Premiums and Runoff (2)	Leverage Risk Variables (3)	Credit Risk Variables (4)	Liquidity Risk Variables (5)	Control Variables (6)
1992–1993	-4.36	8.81	13.92	8.00	4.36
1993–1994	0.64	-1.21	13.24	6.16	12.91
1994–1995	-0.44	3.93	10.48	0.75	19.36
1995–1996	-1.18	4.50	14.36	6.98	27.70
1996–1997	0.13	-0.39	24.65	5.87	16.90
1997–1998	-1.78	1.58	9.62	4.89	22.24
1998–1999	0.19	-0.28	9.93	0.05	16.82
1999–2000	1.98	3.43	16.44	-1.30	20.40
2000–2001	0.43	4.89	18.28	1.12	19.08
2001–2002	-0.94	1.15	18.47	2.73	13.59
2002–2003	0.15	6.52	16.22	1.28	14.12
2003–2004	1.15	5.64	22.18	1.90	12.64
2004–2005	0.63	5.79	17.33	1.36	19.62
Mean	**-0.26**	**5.20**	**15.78**	**3.06**	**16.90**

Table 8 (Continued) Do Jumbo CD Default Premiums or Runoff Enrich the CAMELS-Downgrade Model?

Panel B: Percentage Change in QPS

Out-of-Sample Window (1)	Default Premiums and Runoff (2)	Leverage Risk Variables (3)	Credit Risk Variables (4)	Liquidity Risk Variables (5)	Control Variables (6)
1992–1993	-1.06	3.65	6.25	4.55	-1.70
1993–1994	0.43	6.45	9.31	1.43	-0.72
1994–1995	0.20	3.59	5.39	0.60	1.20
1995–1996	0.00	0.23	3.85	0.68	2.49
1996–1997	-0.20	1.43	3.46	0.61	1.02
1997–1998	-0.17	0.86	2.06	1.03	3.26
1998–1999	0.00	0.40	2.53	0.40	2.39
1999–2000	0.00	0.91	2.62	0.57	2.50
2000–2001	-0.21	0.84	3.45	0.31	2.72
2001–2002	0.66	1.98	5.84	0.85	2.64
2002–2003	-0.38	2.67	2.96	0.29	2.29
2003–2004	0.72	0.96	5.97	-1.55	0.72
2004–2005	0.82	2.45	2.45	-0.49	1.96
Mean	**0.06**	**2.03**	**4.32**	**0.71**	**1.60**

Notes: This table provides alternative measures of the contribution of simple default premiums and runoff to the CAMELS-downgrade model. Column 2 of Panel A shows the impact on power curve areas of removing the two jumbo CD series from the enhanced downgrade model (baseline model plus premiums and runoff). Column 2 of Panel B notes the impact of removing these series on quadratic probability score (QPS). Changes in the QPS and power curve areas are expressed in percentage-change terms to permit direct comparisons. Positive percentage changes for QPS or power curve areas imply that removing the variable block weakens model performance. To facilitate interpretation of changes, columns 3 through 6 show the impact of removing other variable blocks from the CAMELS-downgrade model, such as the control variables (asset size, dummy for CAMELS rating, and dummy for management rating of 2). The evidence suggests default premiums and runoff add nothing to the CAMELS-downgrade model.

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