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Changing the Rules: State Mortgage Foreclosure Moratoria During the Great Depression

David C. Wheelock

Many U.S. states imposed temporary moratoria on farm and nonfarm residential mortgage foreclosures during the Great Depression. This article describes the conditions that led some states to impose these moratoria and other mortgage relief during the Depression and discusses the economic effects. Moratoria were more common in states with large farm populations (as a percentage of total state population) and high farm mortgage foreclosure rates, although nonfarm mortgage distress appears to help explain why a few states with relatively low farm foreclosure rates also imposed moratoria. The moratoria reduced farm foreclosure rates in the short run, but they also appear to have reduced the supply of loans and made credit more expensive for subsequent borrowers. The evidence from the Great Depression demonstrates how government actions to reduce foreclosures can impose costs that should be weighed against potential benefits. (JEL E44, G21, G28, N12, N22)

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Nearly 1 percent of U.S. home mortgages entered foreclosure during the first quarter of 2008, and almost 2.5 percent of all home mortgages were in foreclosure at the end of the quarter.[1] The high number of home mortgages in foreclosure or at risk of foreclosure has prompted calls for government action. On July 30, 2008, President Bush signed the Housing and Economic Recovery Act of 2008 (H.R. 3221), which, among other provisions, included a $300 billion increase in Federal Housing Administration (FHA) loan guarantees to encourage lenders to refinance delinquent home mortgages. Congress also has considered, among other proposals, directing the Federal National Mortgage Association (Fannie Mae) and the Federal Home Loan Mortgage Association (Freddie Mac) to refinance subprime mortgages, and creating a new federal agency to acquire and refinance delinquent mortgages.[2]

The creation of a new federal agency to purchase delinquent mortgages would mimic a similar agency, the Home Owners' Loan Corporation, which was established to refinance delinquent mortgages during the Great Depression. Mortgage delinquency rates rose sharply during the Depression. By one estimate, approximately half of all U.S. urban home mortgages were delinquent as of January 1, 1934 (Bridewell, 1938, p. 172). The Home Owners' Loan Corporation was established in 1933 and over the subsequent three years purchased and refinanced more than 1 million delinquent home loans. Additional steps by the

1 The stock of mortgages in foreclosure during a given quarter includes mortgages that entered foreclosure during that quarter and foreclosures that began in previous quarters that have not yet been completed. These data are from the Mortgage Bankers Association (Haver Analytics).

2 Fannie Mae and Freddie Mac are the two main government-sponsored enterprises that purchase and securitize home mortgages.

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federal government to ease mortgage market pressures during the 1930s included the creation of the Federal Home Loan Bank System to mobilize funds for home lending, the introduction of FHA mortgage insurance, and the creation of Fannie Mae to purchase FHA-insured loans.[3]

State and local governments also responded to the rise in mortgage foreclosures during the Depression, mainly by changing state laws governing foreclosure. Several states enacted temporary foreclosure moratoria. Others made permanent changes that limited the rights or incentives of lenders to foreclose on mortgaged property. Recently a number of U.S. states have considered similar steps to reduce mortgage foreclosures. During the first six months of 2008, the state legislatures of Massachusetts, Minnesota, and New York considered legislation to impose moratoria on foreclosures, and legislation for a national moratorium was introduced in the U.S. Congress.

Foreclosure moratoria are controversial. Although moratoria can benefit some borrowers and temporarily reduce foreclosures, critics argue that moratoria reduce the supply of loans and increase costs for future borrowers.[4] Despite similar arguments made during the Great Depression, 27 states imposed moratoria at the time to reduce the number of mortgage foreclosures.[5] Today, the growing sentiment for using moratoria to reduce the current number of foreclosures prompts a retrospective look at other episodes, such as the Great Depression, when moratoria were used to limit mortgage foreclosures. This article summarizes the main types of mortgage foreclosure laws enacted by U.S. states during the 1930s. Further, it examines why some states elected to impose foreclosure moratoria but others did not. Finally, it summarizes empirical evidence on the costs of foreclosure moratoria borne by borrowers.

MORTGAGE DISTRESS DURING THE GREAT DEPRESSION

The Great Depression was a cataclysmic event. Between 1929 and 1933, U.S. personal income declined 44 percent, real output fell by 30 percent, and the unemployment rate climbed to 25 percent of the labor force. U.S. real estate markets were already showing signs of distress before the Great Depression began. The number of nonfarm residential real estate foreclosures doubled between 1926 and 1929. With the onset of the Depression, the number of foreclosures rose still higher, from 134,900 in 1929 to 252,400 in 1933.[6] The foreclosure rate, shown in Figure 1, increased from 3.6 per 1,000 home mortgages in 1926, the first year data are available, to a high of 13.3 per 1,000 mortgages in 1933. In that year, on average 1,000 home mortgages were foreclosed every day (Federal Home Loan Bank Board, 1937, p. 4). Many more homes were at risk of foreclosure—as many as half of urban home mortgages were delinquent on January 1, 1934 (Bridewell, 1938, p. 172).

The Great Depression also sharply increased farm mortgage foreclosures, which were unusually high throughout the 1920s and 1930s; an average of more than 100,000 farm mortgages entered foreclosure each year from 1926 to 1940. Figure 2 shows that the farm foreclosure rate was especially high from 1932 through 1934, peaking at nearly 39 foreclosures per 1,000 farms in 1933.[7]

The sharp increase in mortgage distress during the Great Depression was the result of precipitous declines in income and real estate values following a period of rapid growth in mortgage debt outstanding.[8] A rising level of debt does

[3] These and other federal government responses to mortgage distress during the Great Depression are described in Wheelock (2008).

[4] For example, see Sloan (2008).

[5] The federal government also enacted a moratorium on farm mortgage foreclosures during the Great Depression. The Frazier-Lemke Farm Bankruptcy Act of 1934 authorized federal courts to grant a five-year moratorium on foreclosure and to scale down a farmer's debt to the current value of his property. The act was declared unconstitutional by the Supreme Court in 1935. Subsequently, Congress enacted the Frazier-Lemke Farm Mortgage Moratorium Act of 1935, which modified and limited the terms of the moratorium. The constitutionality of the latter act was upheld by the Supreme Court in 1937.

[6] *Historical Statistics of the United States, Earliest Times to the Present: Millennial Edition* (2006), series Dc1255.

[7] Alston (1983, Table 1) reports average annual foreclosure rates of 3.2 per 1,000 farms for 1913-20, 10.7 for 1921-25, 19.8 for 1926-40, 3.2 for 1941-50, 1.7 for 1951-60, 1.3 for 1961-70, and 1.3 for 1971-80.

[8] See Alston (1983) and Wheelock (2008) for discussion on the growth of farm and nonfarm mortgage debt, respectively, during the 1910s and 1920s.

Figure 1

Nonfarm Real Estate Mortgage Foreclosure Rate, 1926–1941

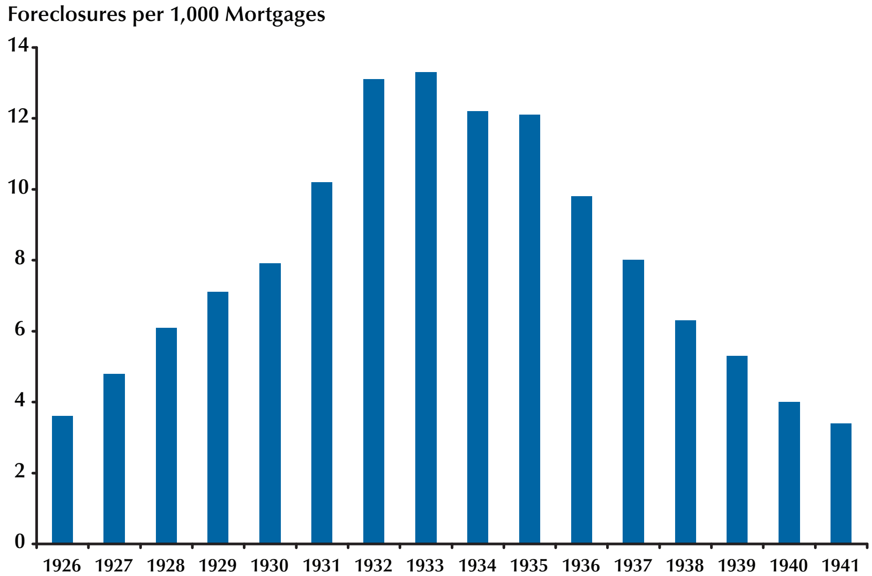


Figure 2

U.S. Farm Foreclosure Rate, 1926–1941

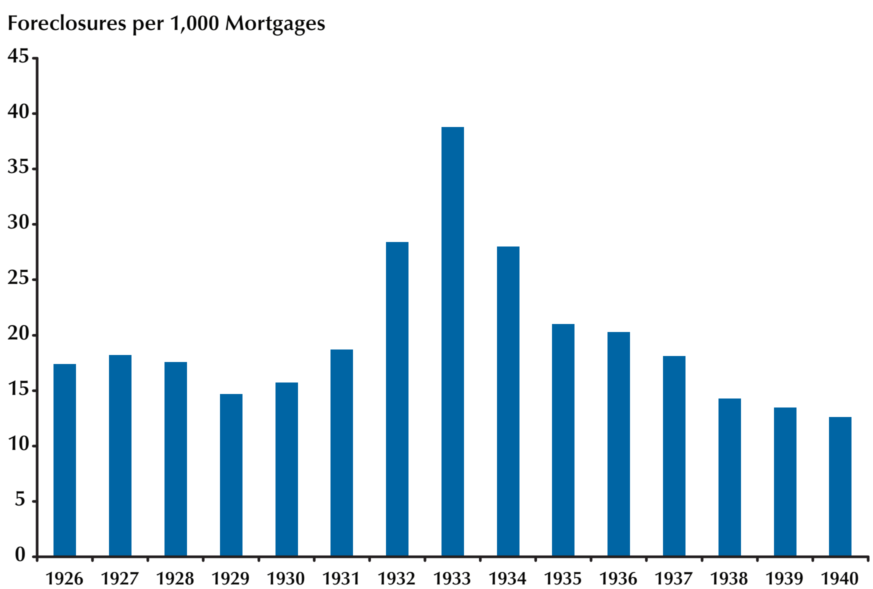
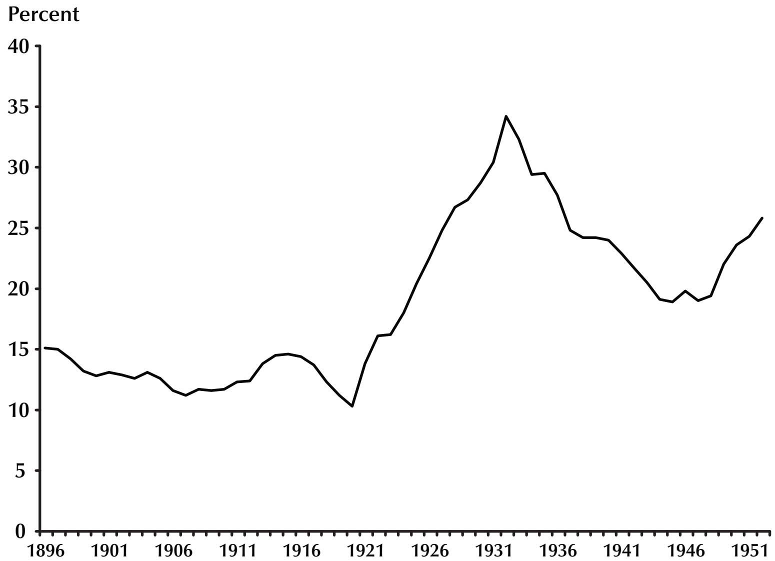


Figure 3

Nonfarm Residential Mortgage Debt as a Percentage of Nonfarm Residential Wealth



SOURCE: Grebler, Blank, Winnick (1956, table L-6).

not necessarily pose a problem for borrowers, provided their incomes and wealth are sufficient to make loan payments. However, between 1929 and 1932, personal disposable income and nonfarm residential wealth fell 41.0 percent and 25.7 percent, respectively, whereas the value of nonfarm residential debt fell a mere 6.8 percent. As shown in Figure 3, nonfarm residential mortgage debt increased sharply relative to nonfarm residential wealth during the 1920s and continued to rise until 1932. Moreover, falling house prices meant that homeowners who were having difficulty making their mortgage payments were increasingly unlikely to sell their homes for more than the outstanding balances on their loans.

Moreover, many home mortgages were short-term, nonamortizing loans that typically were refinanced on maturity.[9] Refinancing usually was easily accomplished during the 1920s, when household incomes and property values were generally rising, but next to impossible during the Depression. Falling incomes made it increasingly difficult for borrowers to make loan payments or to refinance outstanding loans as they came due. The failure of thousands of banks and other lenders made refinancing difficult even for good borrowers; customer relationships were severed and the costs of credit intermediation rose (Bernanke, 1983). The mix of falling household incomes and property values and short-term, nonamortizing loans resulted in soaring mortgage delinquency and foreclosure rates.[10]

Farmers faced similar problems. U.S. farm income fell from $6.2 billion in 1929 to $2.0 billion in 1932. At the same time, farm mortgage

[9] Mortgage lending terms varied considerably across lenders. Savings and loan associations typically made long-term, amortizing mortgage loans. However, banks and life insurance companies often made short-term, nonamortizing (or only partly amortizing) loans. See Morton (1956) for more information about the mortgage market and loan characteristics during the 1920s and 1930s.

[10] As discussed in Wheelock (2008), federal agencies created during the 1930s to rescue and reform the mortgage market encouraged the use of long-term, amortizing mortgage loans—so-called conventional loans. Nonamortizing, unconventional loans have become more common in recent years, however, which some analysts contend has contributed to the increase in mortgage loan delinquencies and foreclosures since 2006.

debt outstanding rose from 40 percent of the value of farm land and buildings in 1930 to 50 percent in 1935.[11] Hence, sharply falling incomes made it increasingly difficult for farmers to pay the interest and principal on their outstanding debts, but falling property values made it less likely that farmers could sell their properties for more than the outstanding balance on their mortgages. The result was a sharp increase in farm mortgage delinquencies and foreclosures.

FORECLOSURE RELIEF LEGISLATION

The first attempts to reduce foreclosures during the Great Depression focused on encouraging lenders and borrowers to renegotiate loan terms through mediation boards and other voluntary arrangements. However, the clamor for compulsory foreclosure moratoria grew louder as the Depression worsened and the number of foreclosures rose. On February 8, 1933, Iowa became the first state to enact a moratorium on mortgage foreclosures. Over the subsequent 18 months, a total of 27 states enacted legislation to limit or halt foreclosures (Skilton, 1944, p. 78).

Mortgage Law

Mortgages and similar loan contracts often are used to finance the purchase of homes, farms, and other real estate.[12] The mortgage contract specifies the terms under which the borrower is obligated to make regular payments of principal and interest to retire the loan. If at some point the borrower fails to make the contracted payments, the loan agreement and laws of the state in which the property is located determine the actions the lender may take to enforce the loan contract. The mortgaged property serves as the security or collateral for the loan, and if the borrower defaults on the mortgage contract, the lender may foreclose on the property against which the loan was made, subject to the state's laws governing foreclosure.

State laws governing the foreclosure process vary. For example, foreclosure may be *judicial* or *nonjudicial*. Under judicial foreclosure, the lender sues the delinquent borrower in court for nonperformance. Typically, judicial foreclosure results in the public sale of the mortgaged property under court supervision, with the proceeds used to satisfy the outstanding mortgage balance and any other outstanding liens on the property.

Under nonjudicial foreclosure by "power of sale," the mortgaged property is sold without court supervision in the event of borrower default, again with the sale proceeds used to pay the outstanding balance of the mortgage and any other liens. Some states permit *strict foreclosure*, which grants the lender unconditional title to the mortgaged property in the event of borrower default.

The laws of some states grant statutory redemption periods during which a borrower (mortgagor) may regain ownership of a property after foreclosure sale by payment of the foreclosure sale price, interest, and taxes. Generally, redemption is permitted from six months to one year after the foreclosure sale. During the Depression, several states modified their laws to extend or enhance the rights of mortgagors to redeem foreclosed property.

Finally, some states allow *deficiency judgments* in which a mortgage holder is granted a lien against other assets of the borrower when the proceeds from a sale of the mortgaged property do not cover the outstanding mortgage balance.[13] During the Depression, several states enacted reforms that limited the rights of lenders to seek deficiency judgments against borrowers.

Examples from the Great Depression

The diversity of foreclosure proceedings across U.S. states during the 1930s was noted in a 1936 federal government study:

[11] *Historical Statistics of the United States, Earliest Times to the Present: Millennial Edition* (2006), series Da1295 (farm income) and series Da579 (debt as a percentage of land and building value).

[12] Deeds of trust are used to finance real estate purchases in some states. Unlike a mortgage, a deed of trust involves an independent trustee who holds a power of sale in the event of default and who conveys the property to the borrower once the deed of trust is paid in full. See McDonald and Thornton (2008) for basic information about the mortgage market and mortgage finance.

[13] If the value of a property exceeds the outstanding loan balance, the borrower generally benefits from refinancing the loan or selling the property and paying off the outstanding loan balance rather than losing the property through foreclosure. Hence, the proceeds from most foreclosure sales are less than the outstanding mortgage balance.

> A general survey indicates that in twenty-eight states foreclosure is by action in court. Ten states use unregulated power of sale. Five states use regulated power of sale, and the remaining states have various other methods. Thirty-one states provide a redemption period ranging from four months in Oregon to two years in Alabama. Seventeen states have no redemption period, but, of these, eight use foreclosure in court which requires months to complete. (Central Housing Committee, 1936, p. 2)

During the Depression, many states enhanced borrower redemption rights, limited deficiency judgments, or made other changes that favored borrowers, and several states imposed moratoria on foreclosures. The specific details of moratoria legislation varied widely. A few states imposed blanket moratoria that temporarily prohibited most foreclosures of farm and nonfarm home mortgages contracted before a specified date. However, most states limited their moratoria to specific situations. For example, some states granted relief only for borrowers who were current in the payment of interest and taxes but delinquent in the payment of loan principal. For example, a New York statute enacted in 1933 specified that "No action for the foreclosure of a mortgage on real estate solely on account of default in payment of principal…shall be brought before July 1, 1937" (Central Housing Committee, 1936, p. A-18). Foreclosures were permitted, however, against borrowers who had ceased to pay interest and taxes, as well as principal.

Several states directed their state courts to grant moratoria in deserving cases, but little guidance was provided to the courts about how to determine which borrowers deserved relief. For example, in Iowa, the court was authorized to grant a borrower's request for relief from pending foreclosure unless "good cause is shown to the contrary" (Skilton, 1944, p. 82). Similarly, an Arizona statute specified that "In pending or future real estate mortgage foreclosure suits, the court may order a two-year continuance unless good cause to the contrary is shown" (Central Housing Committee, 1936, p. A-3). Not surprisingly, the extent to which courts granted relief to delinquent borrowers varied widely, even within a state. Many courts determined that it was pointless to grant relief to borrowers who had no hope of refinancing their mortgage or making payments or who did not act in good faith toward their lender (Skilton, 1944, pp. 98-106). In addition, courts often required borrowers to pay rent or interest to the lender, as well as taxes, as a condition for halting foreclosure proceedings.

In conjunction with a foreclosure moratorium, several states extended the period during which a mortgagor could redeem his property after foreclosure. Again, however, any extension of the redemption period was often left to the court's discretion. In Kansas, for example, "the period for redemption on real estate may be extended for such additional time as the court shall deem it just and equitable" (Central Housing Committee, 1936, p. A-10). In a few states, the legislation was more specific. For example, North Dakota legislation specified that "The period within which a mortgagor or judgment debtor may redeem from a mortgage foreclosure or execution sale of real estate…is extended for a period of two years" (Central Housing Committee, 1936, p. A-21).

Several states also modified their statutes to limit deficiency judgments. Some states restricted judgments to the difference between the outstanding loan balance and a "fair" or "reasonable" value of the mortgaged property, rather than the difference between the loan balance and the price received at a foreclosure sale. For example, a 1933 Idaho statute specified that "no deficiency judgment may be entered in any amount greater than the difference between the mortgage indebtedness, plus the cost of foreclosure and sale and the reasonable value of the property" (Central Housing Committee, 1936, p. A-7). Other states permitted courts to invalidate foreclosure sales for less than fair value. Most states left the determination of fair value to the discretion of a local appraisal board or court rather than attempt to define "fair value" in statutes.

Several states imposed new limits on the length of time that a lender could seek a deficiency judgment after a foreclosure sale. For example, Iowa and Ohio enacted legislation limiting deficiency judgments to two years after a foreclosure sale (Skilton, 1944, p. 130). Other states abolished the right of lenders to seek deficiency judgments altogether. For example, a 1935 Montana statute specified that "Deficiency judgments are abolished in all actions for foreclosure of mortgages for balance of purchase price of real property" (Central Housing Committee, 1936, p. A-16).[14]

WHICH STATES ADOPTED MORATORIA AND WHY?

The 27 states that adopted foreclosure moratoria during 1933 and 1934 are listed in Table 1, and the geographic distribution of states with moratoria is shown in Figure 4. Moratoria were especially common among states in the Midwest and Great Plains, but they also were imposed by several states in the Northeast and Far West. Foreclosure moratoria were less common in New England, the Southeast, and Mountain West.[15]

Foreclosure moratoria generally applied to both farm and nonfarm residential mortgages. However, the pressure for foreclosure moratoria was particularly intense in midwestern states where farm foreclosure rates were especially high (Figure 5). Moratoria were less common in states with relatively low farm foreclosure rates, though a few, including New Hampshire, Pennsylvania, and Vermont, also imposed moratoria.

Alston (1984) investigates why some, but not all, states imposed foreclosure moratoria during the Depression. He estimates a logit regression model that includes a state's farm foreclosure rate, percentage of farms mortgaged, and percentage of farm mortgages held by federal land banks as explanatory variables. Alston argues that a state was more likely to impose a moratorium the higher its farm foreclosure rate, the higher its percentage of mortgaged farms, and the lower the percentage of mortgages held by federal land banks (which were less likely to foreclose than other lenders).[16] He finds that the farm foreclosure rate had the strongest impact on a state's decision to impose a moratorium.

As noted previously, moratoria were adopted in a few states with relatively low farm foreclosure

Table 1

State Mortgage Moratoria during the Great Depression

States imposing moratoria	States not imposing moratoria
Arizona	Alabama
Arkansas	Colorado
California	Connecticut
Delaware	Florida
Idaho	Georgia
Illinois	Indiana
Iowa	Kentucky
Kansas	Maine
Louisiana	Maryland
Michigan	Massachusetts
Minnesota	Missouri
Mississippi	New Jersey
Montana	New Mexico
Nebraska	Nevada
New Hampshire	Rhode Island
New York	Tennessee
North Carolina	Utah
North Dakota	Virginia
Ohio	Washington
Oklahoma	West Virginia
Oregon	Wyoming
Pennsylvania	
South Carolina	
South Dakota	
Texas	
Vermont	
Wisconsin	

SOURCE: Skilton (1944, p. 78).

[14] See Central Housing Committee (1936), Poteat (1938), or Skilton (1944) for additional information about the provisions of moratoria and other legislation affecting the rights of mortgagors and lenders enacted in different states during the Depression.

[15] The source for Table 1 and Figure 4 is Skilton (1944, p. 78), which lists 27 states as having had a moratorium. Other sources omit Oregon, where a moratorium was authorized by a joint resolution of the state legislature, rather than by statute (Poteat, 1938), or omit both Oregon and Arkansas (Alston, 1984).

[16] The Federal Farm Loan Act of 1916 established 12 regional federal land banks to increase the supply of farm mortgage loans. See www.fca.gov/about/history/historyFCA_FCS.html.

Figure 4

States Imposing Foreclosure Moratoria During the Great Depression

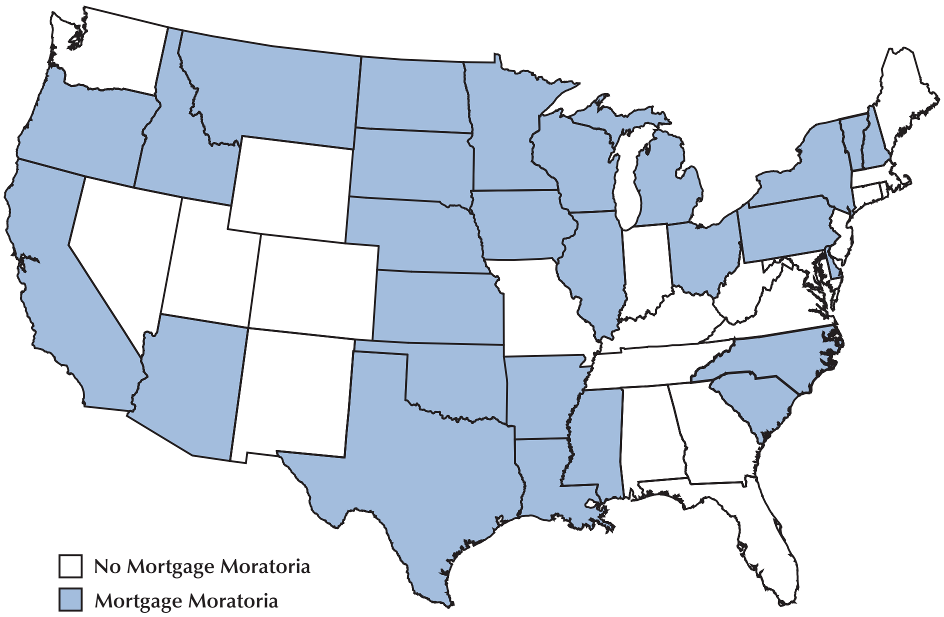


Figure 5

Average Farm Foreclosure Rates, 1929–1932

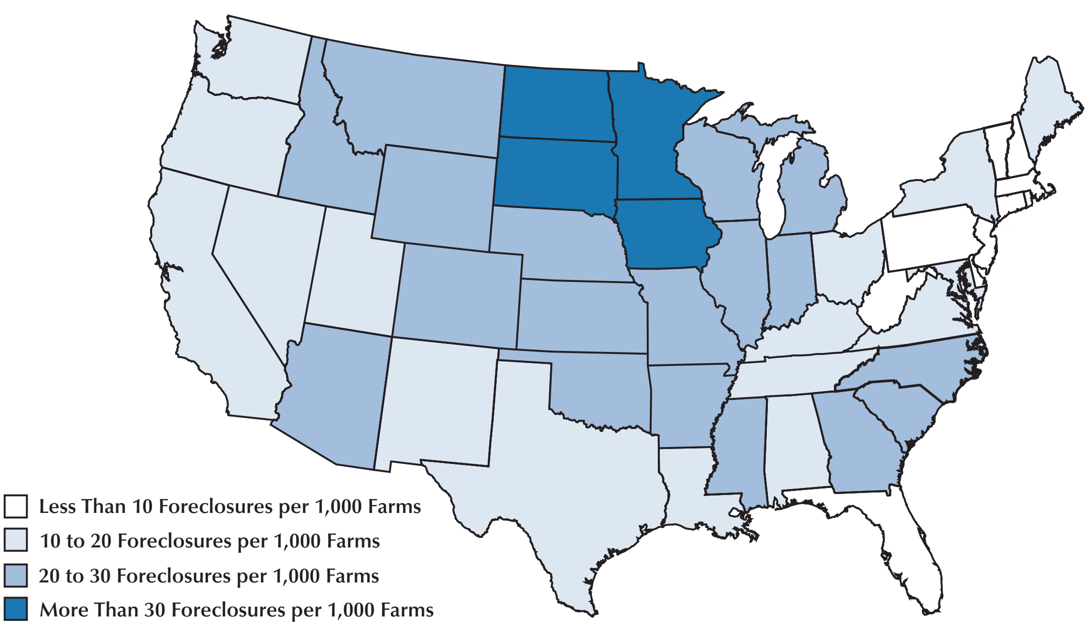


Table 2

Regression Results

	Model			
Variable	1	2	3	4
Intercept	–1.7768	–4.0343	–1.5403	–2.4317
	(1.7752)	(3.4049)	(2.9175)	(3.6759)
Farm foreclosure rate	**0.0803****	**0.1338***	–0.0183	0.0419
	(0.0366)	(0.0695)	(0.0647)	(0.1081)
Mortgaged farm percent	2.0338	3.2953	1.7109	2.033
	(3.4587)	(4.3798)	(3.7252)	(4.5092)
Federally held farm debt	–3.0984	–3.4411	**–6.4965***	–5.5301
	(2.5978)	(–3.1007)	(3.7163)	(3.8831)
Owner-occupied nonfarm homes		3.5653	2.9328	3.6291
		(5.9059)	(5.5364)	(5.9293)
Foreclosure rate × farm population			**0.0023***	0.0017
			(0.0013)	(0.0016)
Midwest		–1.7877		–1.0099
		(1.8866)		(2.0200)
South		–0.5345		–0.6759
		(1.4734)		(1.4776)
West		–2.0764		–1.3836
		(1.4546)		(1.5882)
Log likelihood	–27.3815	–25.8424	–25.6897	–25.2482
Probability > chi-square	0.0116	0.0493	0.0132	0.0537

NOTE: Standard errors are indicated in parentheses; statistically significant coefficients are in bold. *Indicates significance at the 90 percent confidence level; ** indicates significance at the 95 percent confidence level.

See the Appendix for data definitions and sources.

rates, and some states with high farm foreclosure rates did not impose moratoria. According to Skilton (1944), some states imposed moratoria in response to high numbers of nonfarm home mortgage foreclosures. Unfortunately, state-level data on nonfarm real estate foreclosures are not available for the early 1930s to test directly the impact of nonfarm foreclosures on moratoria adoption. Nevertheless, regional differences in farm foreclosure rates and the adoption of moratoria suggest that nonfarm foreclosures or other considerations may have influenced the decision to impose moratoria in some states.

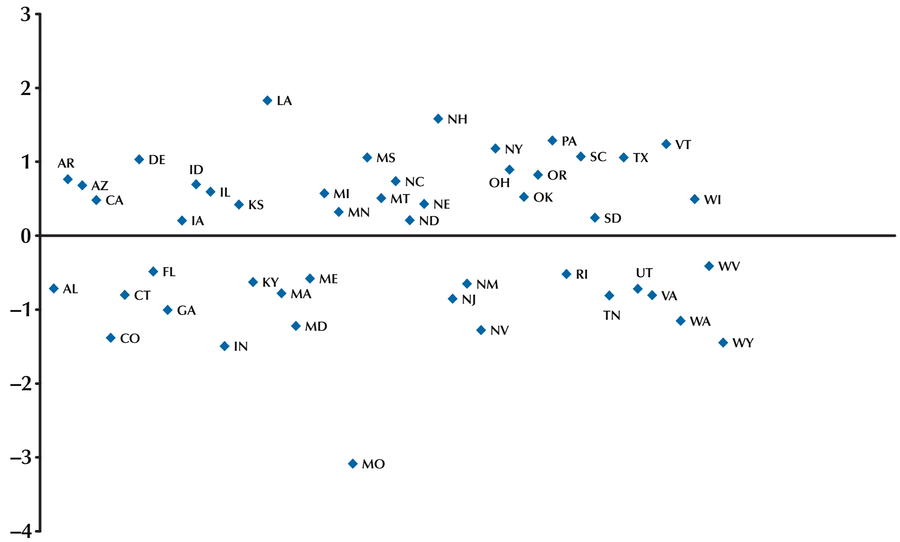
Some evidence on why states imposed foreclosure moratoria is reported in Table 2, which presents a replication of Alston's (1984) logit model and some alternative specifications. The dependent variable in this set of cross-sectional regressions is a dummy variable, set equal to 1 for states that adopted a moratorium during 1933-34 and to 0 otherwise. The Appendix provides complete definitions and data sources for the variables included in the regressions.

Model 1 replicates Alston's model and shows his main result: the higher a state's farm foreclosure rate, the greater the likelihood the state would adopt a foreclosure moratorium.[17] Model 2 includes the percentage of owner-occupied nonfarm homes as an additional explanatory variable.

[17] The coefficient estimates in Model 1 differ slightly from those reported in Alston (1984). Unlike Alston, I treated Arkansas and Oregon as having had moratoria, based on Skilton (1944), and used the farm foreclosure rate for 1932, rather than the average farm foreclosure rate for 1932 and 1933, as an explanatory variable.

Figure 6

Model 1: Pearson Residuals



Presumably, the demand for moratoria legislation was greater where a high percentage of homes were mortgaged and, hence, at risk of foreclosure. Unfortunately, state-level data on the percentage of homes carrying a mortgage are not available for the 1930s. However, if owner-occupied homes were no less likely to be mortgaged than rented homes, a higher percentage of owner-occupied homes might reflect a greater demand for a foreclosure moratorium. Accordingly, I expect a positive coefficient on this variable. Consistent with expectations, the coefficient estimate in Model 2 for the percentage of owner-occupied homes is positive, though not statistically significant.

Model 2 also includes regional dummy variables. The coefficient estimates for the regional dummies indicate that relative to the Northeast (the omitted region), states in other regions of the country were less likely to adopt foreclosure moratoria. Stated differently, for a given rate of farm foreclosures, states in the Northeast were more likely to adopt foreclosure moratoria than states elsewhere. This suggests that nonfarm mortgage distress had a greater influence on the decision to adopt moratoria among the more urbanized northeastern states than it did in other regions of the country. However, the coefficients on the regional dummy variables are not statistically significant.[18]

Model 3 further refines the analysis by testing whether the influence of farm foreclosures on moratoria adoption was stronger in states with relatively high farm populations. To test this conjecture, Model 3 includes the interaction of the farm foreclosure rate and the percentage of state population located on farms. The coefficient estimate on the interaction term (*foreclosure rate* × *farm population*) is positive and statistically significant, and the coefficient on the farm foreclosure rate is near zero, which supports this hypothesis. The impact of a given farm foreclosure rate was greater in states with relatively larger farm populations.

Finally, Model 4 adds regional dummy variables to the previous specification. The coefficients on the regional dummies are again negative, suggesting a relatively low demand for moratoria

[18] The test statistic for a likelihood ratio test of the hypothesis that the coefficients on the regional dummies are jointly zero is 3.08 (p-value = 0.38).

among states outside the Northeast. However, the contribution of the regional dummies to the model's explanatory power is not statistically significant.[19]

For additional insights about why states adopted (or did not adopt) foreclosure moratoria, I examined the residuals from the logit models reported in Table 2. The Pearson residuals for Model 1 are shown in Figure 6.[20] The residuals for states that adopted moratoria are greater than or equal to zero, whereas those for states that did not adopt moratoria are less than or equal to zero. The closer a state's residual is to zero, the more accurately the model explains the state's decision to impose (or not impose) a moratorium. Thus, the large positive residual for Louisiana indicates that the model explains relatively little of the state's decision to impose a moratorium. Similarly, the model explains relatively little of Missouri's decision not to adopt a moratorium. Missouri had a comparatively high farm foreclosure rate and, given that fact, Model 1 predicts that Missouri would have imposed a moratorium.

Additional information helps explain the anomalous behavior of some states. For example, Louisiana was the last state to adopt a debt moratorium (in July 1934). Soon thereafter, the legislation authorizing the moratorium was amended to grant broad authority to a state debt commissioner to "suspend all laws relating to the collection of fundamentally all types of debts in existence at the time of the passage of the act" (Skilton, 1944, pp. 83-84). In effect, the state imposed a general moratorium on all debts, not just real estate mortgage debt. The breadth of the moratorium thus might help explain why Louisiana imposed a moratorium, despite only a modest level of farm distress.

New York also enacted an unusually broad foreclosure moratorium that extended to commercial real estate, as well as to farm and nonfarm residential property. According to Skilton (1944, pp. 76-77), property management companies and other real estate interests had considerable influence on the moratorium legislation, and "The lobbying of real estate operators was sufficient… to defeat Governor Lehman's original idea that a moratorium should be limited to farms and homes." Thus, like Louisiana, the breadth of the moratorium may help explain why Model 1, which captures only the effects of farm distress, does not explain well the imposition of a foreclosure moratorium in New York.

Differences in the prevailing state laws governing mortgage foreclosure might also help account for the model's failure to explain well the moratoria decisions of some states. For example, neither Indiana nor Missouri adopted a foreclosure moratorium during the Depression, despite relatively high levels of farm distress. However, a federal study concluded that the demand for moratoria was low in both states because their prevailing foreclosure laws were already comparatively favorable to borrowers (Central Housing Committee, 1936).

ECONOMIC IMPACT OF FORECLOSURE MORATORIA

Governments cause both immediate and long-term effects when they rewrite the terms of contracts between private parties. The immediate impact is redistribution of wealth between the parties of the affected contracts. The temporary foreclosure moratoria and most other changes in state mortgage laws enacted during the 1930s favored borrowers over lenders. These actions interfered with the rights of lenders to seize collateral pledged by borrowers to guarantee payment of their mortgages. Several states also enhanced the rights of borrowers to redeem foreclosed property and limited the rights of lenders to sue for deficiency judgments.

One immediate effect of mortgage relief legislation during the Depression was reduced farm foreclosure rates (Rucker and Alston, 1987).[21]

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[19] The test statistic for a likelihood ratio test of the hypothesis that the coefficients on the regional dummies are jointly zero is 0.88 (p-value = 0.83).

[20] The logit models were estimated using Stata/MP 10.0. The basic Pearson residual is the difference between the actual and model-predicted values of the dependent variable, divided by the estimated standard deviation of the predicted values. See Stata/MP 10.0 for more details about the calculation of the Pearson residual.

[21] I am unaware of any research on the effects of relief legislation on nonfarm home mortgage foreclosure rates.

However, over the longer run, foreclosure moratoria and other changes in mortgage laws may have made loans costlier or more difficult to obtain. Critics argued that foreclosure moratoria induce lenders to restrict the supply of loans and raise interest rates to compensate for the possibility that their right to foreclose on delinquent loans or to collect deficiency judgments will be constrained. According to a 1936 federal government report,

> Statutes which provide a lengthy, expensive, complicated or otherwise burdensome foreclosure procedure, or which interpose a long period of redemption before title and possession to the mortgaged property can be obtained, have a tendency to increase interest rates and security requirements throughout the jurisdiction, since prospective lenders naturally take into account the procedure available for realizing the debt out of the security when determining the conditions on which they will be willing to make loans. (Central Housing Committee, 1936, p. 3)

The same report noted that in 1933-34 many states elected to disregard such objections because it was widely believed that "unrestricted foreclosure of farm and home mortgages under the circumstances prevailing at the time would have deprived large numbers of persons of essential shelter and protection, and would have left them without the necessary means for earning a living. Such wholesale evictions might have seriously endangered basic interests of society" (Central Housing Committee, 1936, p. 2). Hence, in many states, the societal costs of widespread foreclosures were viewed as exceeding the costs of reduced loan supply and higher interest rates borne by prospective borrowers. Furthermore, foreclosure moratoria generally were viewed as expedients to buy time for the economy to recover and for the federal government to initiate programs to refinance delinquent mortgages (Skilton, 1944, pp. 73-77). Even lenders may have benefited from foreclosure moratoria in the short run. Although individual lenders had an incentive to foreclose to recoup losses on delinquent mortgages, a high number of foreclosures in an area could reduce property values and thereby cause still more foreclosures. Thus, foreclosure moratoria might halt a downward spiral in property values and benefit lenders as a whole.[22]

Although the economic and societal benefits of lower foreclosure rates are difficult to measure, research shows that the foreclosure moratoria of the Great Depression did impose costs on future borrowers. Alston (1984) investigates the impact of foreclosure moratoria in an empirical model of the farm mortgage market. He argues that foreclosure moratoria encouraged lenders to reduce the supply of loans, resulting in fewer loans made and, possibly, higher average interest rates. Consistent with this hypothesis, Alston (1984) finds that private lenders made significantly fewer loans in states that imposed moratoria and tended to charge higher interest rates on the loans they did make.

Rucker (1990) extends Alston's (1984) study to investigate differences in the impact of mortgage relief legislation on the supply of loans offered by different types of private lenders. In the 1930s, most farm mortgages were issued by local commercial banks, private individuals, insurance companies, and federal land banks. Insurance companies tended to be larger and more diversified and to have a lower cost of funds than did banks and individual lenders. Their size and cost advantages enabled insurance companies to attract lower-risk borrowers and, consequently, experience lower delinquency rates. Insurance companies generally were also more willing to grant extensions to delinquent borrowers. Hence, the costs imposed by mortgage relief legislation should have been lower for insurance companies than for other private lenders. Rucker (1990) finds that, indeed, mortgage relief legislation led to significantly larger reductions in the supply of loans from commercial banks and individual lenders than from insurance companies.[23] Both

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[22] Kahn and Yavas (1994) examine the short- and long-run effects of changes in foreclosure laws (especially how they affect borrower and lender behavior and borrower welfare) in a simple theoretical model of the mortgage market in which renegotiation of loan contracts is possible. Jaffe and Sharp (1996) describe the economics of foreclosure moratoria in the context of alternative legal theories of contracts.

[23] In his econometric analysis, Rucker (1990) treated legislation that limited deficiency judgments or enhanced redemption rights for

Alston (1984) and Rucker (1990) conclude that mortgage relief legislation caused significant reductions in the aggregate supply of loans in states that enacted such legislation.

The findings of Alston (1984) and Rucker (1990) on the effects of mortgage relief legislation during the 1930s are consistent with other studies that find significant effects of state mortgage laws on local lending markets. Meador (1982), for example, finds that loan interest rates tend to be higher in states with lengthy or costly foreclosure processes or those that prohibit deficiency judgments. More recently, Pence (2006) finds that mortgage loans are, on average, some 3 to 7 percent smaller in states in which foreclosure requires a court action than in states with nonjudicial foreclosure processes, again consistent with the hypothesis that the supply of loans is lower in states in which foreclosure is more costly.[24]

CONCLUSION

In 2008, residential real estate foreclosure rates are at their highest levels since the Great Depression. Not surprisingly, policymakers are considering actions similar to those taken during the Depression to limit foreclosures. The federal government responded to mortgage distress during the Depression by creating new federal agencies to refinance delinquent mortgages, insure and finance newly issued mortgages, and expand federal farm credit programs. By contrast, many state governments imposed moratoria on foreclosures, limited deficiency judgments, and enhanced the rights of borrowers to redeem foreclosed property. By halting foreclosures temporarily, states hoped to buy time for economic recovery to take hold, for household incomes and property values to rise, and for the federal government to refinance delinquent mortgages.

The earliest calls for mortgage relief were in farming regions, and states with high farm foreclosure rates were more likely to impose moratoria (Alston, 1984). Additional evidence indicates that farm foreclosures had a greater impact on the decision to impose moratoria in states in which the farm population comprised a relatively high percentage of total state population.

Moratoria were imposed in a few states with comparatively little farm mortgage distress, suggesting that urban mortgage distress or other factors influenced the decision to impose moratoria in some states. For example, in New York, lobbying by commercial real estate interests helped shape legislation for a broad moratorium covering farm, urban residential, and commercial real estate mortgage foreclosures.

In most states, foreclosure moratoria were limited to borrowers who had some chance of paying or refinancing their loans. Relief often was denied to borrowers judged to have little prospect of ever paying off their mortgage.

Foreclosure moratoria resulted in both winners and losers. Although the rights of lenders to foreclose on collateral or to seek deficiency judgments were restricted, relief legislation did apparently contribute to a reduction in farm failures (Rucker and Alston, 1987).

At least some contemporaries recognize that even temporary foreclosure moratoria can impose costs on future borrowers. Alston (1984) and Rucker (1990) find that lenders reduced the supply of loans in response to diminution of their rights to foreclose on collateral or to seek deficiency judgments. Thus, while to many observers the economic and societal costs of widespread real estate foreclosures were overwhelming, foreclosure moratoria and other relief legislation transferred at least some of those costs to future borrowers. The evidence from the use of foreclosure moratoria during the Great Depression demonstrates how legislative actions to reduce foreclosures can impose costs that should be weighed against potential benefits.

borrowers, as well as foreclosure moratoria, as forms of relief legislation, whereas Alston (1984) focused exclusively on moratoria.

[24] Pence (2006) compares bordering census tracts located in different states and controls for a variety of borrower, policy, and other census tract characteristics.

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APPENDIX

Variable Definitions and Data Source Information

Variable name	Definition	Source
Moratorium	Dummy variable equal to 1 for states with mortgage moratorium in 1933-34	Skilton, Robert H. *Government and the Mortgage Debtor (1929 to 1939)*. PhD Dissertation, University of Pennsylvania, Philadelphia, 1944, p. 78.
Farm foreclosure rate	Farm foreclosures per 1,000 mortgages in 1932	U.S. Department of Agriculture. "The Farm Real Estate Situation, 1930-31." Bureau of Agricultural Economics, circular no. 209, 1931.
Mortgaged farms (percent)	Percentage of farms mortgaged in 1930, calculated as (mortgaged farms/all owned farms)	U.S. Department of Commerce. *Statistical Abstract of the United States*: *1932*. Washington, DC: U.S. Government Printing Office, 1932, Table 548, p. 589.
Federally held farm debt	Percent of mortgage debt held by federal land banks, calculated as (sum of amount of loans closed 1917 to 1932/ total farm mortgage debt in 1932)	U.S. Department of Agriculture. Miscellaneous Publication No. 478, "Farm Mortgage Credit Facilities in the United States." Washington, DC: U.S. Government Printing Office, 1942, Table 64, p. 221 and Table 78, p. 245.
Owner-occupied nonfarm homes (percent)	Percentage of owned nonfarm homes in 1930, calculated as (sum of owned nonfarm homes/total nonfarm homes)	U.S. Department of Commerce. *Fifteenth Census of the United States: 1930*. *Population*, Volume VI, Table 42, p. 35. Washington, DC: U.S. Government Printing Office, 1931.
Farm population	Percentage of population on farms in 1930	U.S. Department of Commerce. *Statistical Abstract of the United States: 1932*. Washington, DC: U.S. Government Printing Office, 1932, Table 36, p. 47.

Mortgage Innovation, Mortgage Choice, and Housing Decisions

Matthew S. Chambers, Carlos Garriga, and Don Schlagenhauf

This paper examines some of the more recent mortgage products now available to borrowers. The authors describe how these products differ across important characteristics, such as the down payment requirement, repayment structure, and amortization schedule. The paper also presents a model with the potential to analyze the implications for various mortgage contracts for individual households, as well as to address many current housing market issues. In this paper, the authors use the model to examine the implications of alternative mortgages for homeownership. The authors use the model to show that interest rate–adjustable mortgages and combo loans can help explain the rise—and fall—in homeownership since 1994. (JEL E2, E6)

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Housing is a big-ticket item in the U.S. economy. At the macro level, residential housing investment accounts for 20 to 25 percent of gross private investment. In the aggregate, this financing is about 8 trillion dollars and uses a sizable fraction of the financial resources of the economy. The importance of housing at the individual household level is more evident because the purchase of a house is the largest single consumer transaction and nearly always requires mortgage financing. This decision affects the overall expenditure patterns and asset allocation decisions of the household.

In recent years, interest in the role of housing in the U.S. economy has increased, influenced mainly by two events. During the economic downturn in 2000, the housing sector seemed to mitigate the slowdown in many other sectors of the economy as residential investment remained at high levels. More recently, the large number of foreclosures has again focused attention on the importance of housing. Fears have increased that mortgage market problems will have long-lasting effects on the credit market and thus continue to create a drag on the economy.

Events illustrating the important role of housing in the economy are not limited to those of the past decade. Housing foreclosures soared during the Great Depression as a result of two factors. The mortgage system was very restrictive: Homeowners were required to make down payments that averaged around 35 percent for loans lasting only five to ten years. At the end of the loan period, mortgage holders had to either pay off the loan or find new financing. The 1929 stock market collapse resulted in numerous bank failures. Mortgage issuance fell drastically, and homeowners were dragged into foreclosure. Faced with these problems, the government developed new housing policies as part of the New Deal legislation. The Home Owners' Loan Corporation (HOLC) and the Federal Housing Administration (FHA)

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were created along with a publicly supported noncommercial housing sector. The HOLC was designed to help distressed homeowners avert foreclosure by buying mortgages near or in foreclosure and replacing them with new mortgages with much longer durations. The HOLC financed these purchases by borrowing from the capital market and the U.S. Treasury. The FHA introduced new types of subsidized mortgage contracts by altering forms and terms, as well as mortgage insurance. In addition, Congress created Federal Home Loan Banks in 1932 and the Federal Home Loan Mortgage Corporation, commonly known as Fannie Mae, in 1938. The latter organization was allowed to purchase long-term mortgage loans from private banks and then bundle and securitize these loans as mortgage-backed securities.[1] These changes had an important impact on the economy: The stock of housing units increased 20 percent during the 1940s, and the homeownership rate increased approximately 20 percentage points from 1945 to 1965.

The need for increased understanding of housing markets, housing finance, and their linkage to the economy—the objective of this paper—should be obvious. We begin by examining the structure of a variety of mortgage contracts. Given the array of available mortgage products, mortgage choice can be a complex problem for potential home buyers. Buyers must consider many dimensions, such as the down payment, maturity of the contract, repayment structure, the ability to refinance the mortgage, and the impact of changes in interest rates and housing prices. We present examples to clarify key features of prominent mortgage contracts. The best mortgage for one household need not be the best mortgage for another. In fact, a model is needed to understand the mortgage decisionmaking process and what the aggregate implications are for the economy. This model must explicitly recognize the differences among households in age, income, and wealth. In addition, these decisions must reflect the complexities of the tax code that favor owner-occupied housing. Such a framework allows individual decisions to be aggregated so that the impact of mortgage decisions for the economy can be clearly identified.

The second part of this paper presents a model for understanding the impact of mortgage decisions on the economy. We use the model to show the role that adjustable-rate mortgages (ARMs) and combo loans have played since 1994 in the rapid rise—and subsequent decline—in homeownership.

MORTGAGE CONTRACTS

A mortgage contract is a loan secured by real property. In real estate markets this debt instrument uses the structure (building) and land as collateral. In most countries mortgage lending is the primary mechanism to finance the acquisition of residential property. Mortgage loans typically are long-term contracts and require periodic payments that can cover interest and principal. Lenders provide the funds to finance the loans. Usually, such loans are sold to secondary market parties interested in receiving an income stream in the form of the borrower's payments.

The financial marketplace offers many types of mortgage loans, which are differentiated by three characteristics: the payment structure, the amortization schedule, and the term (duration) of the mortgage loan. The payment structure defines the amount and frequency of mortgage payments. The amortization schedule determines the amount of principal payments over the life of the mortgage. This schedule differs across types of mortgage loans and can be increasing, decreasing, or constant. Some contracts allow for no amortization of principal and full repayment of principal at a future, specified date. Other contracts allow negative amortization, usually in the initial years of the loan.[2] The term or duration usually refers to the maximum length of time allotted to repay the

[1] This increased the flow of resources available in areas in which savings were relatively scarce. The intent was to increase the opportunities for low-income families in the housing market. Because of the implicit backing of the government, the riskiness of these assets was perceived to be similar to the risk of U.S. Treasury securities.

[2] A mortgage contract with negative amortization means the monthly payment does not cover the interest on the outstanding balance. As a result, the principal owed actually increases. We illustrate such a contract later in the paper.

Table 1

Types of Primary Mortgage Contracts*

Type of contract	1993*	1995	1997	1999	2003	2005
Fixed-rate mortgage	84.4	82.6	86.5	90.6	92.8	90.0
Adjustable-rate mortgage	11.0	12.3	9.3	5.9	4.3	5.9
Adjustable-term mortgage	0.2	0.0	0.8	0.8	0.3	0.4
Graduated-payment mortgage	1.0	1.0	1.2	1.0	0.7	1.2
Balloon	0.9	1.6	1.0	0.9	1.1	1.2
Other	1.7	1.6	0.1	0.0	0.1	0.1
Combination of the above	0.8	0.9	1.1	0.8	0.7	1.1
Sample size	37,183	39,026	35,999	39,034	42,411	45,450

NOTE: *Share of total contracts in percent.

SOURCE: U.S. Department of Commerce, *American Housing Survey* for various years.

mortgage loan. The most common mortgage contracts are for 15 and 30 years. The combination of these three factors allows a large variety of distinct mortgage products.

Mortgage contracts affect consumer decisions. For example, some contracts are more effective in allowing increased homeownership for younger households. What types of mortgage contracts are actually held in the United States? According to the 2001 *Residential Finance Survey* (U.S. Census Bureau, 2001), roughly 97 percent of housing units were purchased through mortgage loans, whereas only 1.6 percent were purchased with cash. Table 1 summarizes the types of mortgage contracts used in the United States. The fixed-rate (payment) mortgage loan is the dominant contract, and the popularity of an adjustable (or floating) rate mortgage is substantially smaller. In contrast, in the United Kingdom and Spain, where the homeownership rate is 71 and 80 percent, respectively, the adjustable (or floating) rate contract is the dominant contract. The popularity of the fixed-rate contract in the United States is largely a result of the policies of the FHA, Veterans Administration, and various government incentives to sell the loan in the secondary market. This is the role of enterprises such as Fannie Mae and the Federal Home Loan Mortgage Corporation (Freddie Mac), two government-sponsored enterprises (GSEs) that are among the largest firms that securitize mortgages. Mortgage securitization occurs when a mortgage contract is resold in the secondary market as a mortgage-backed security. In the early 1990s, substantial changes occurred in the structure of the mortgage market in the United States. According to data in the 2007 *Mortgage Market Statistical Annual*, the market share of nontraditional mortgage contracts has increased since 2000. Nontraditional or alternative mortgage products include interest-only loans, option ARMs, loans that couple extended amortization with balloon-payment requirements, and other contracts of alternative lending. For example, in 2004 these products accounted for 12.5 percent of origination loans. By 2006, this segment increased to 32.1 percent of loan originations. Given the declining share of conventional and conforming loans, the structure of mortgage contracts merits further consideration.

General Structure of Mortgage Contracts

Despite all their differences, mortgage loans are just special cases of a general representation. Some form of notation is needed to characterize this representation. Consider the expenditure associated with the purchase of a house of size h and a unit price of p. We can consider h as the number of square feet in the house and p as the price per square foot. If buyers purchase a house with cash, the total expenditure is then denoted by ph. Most buyers do not have assets available that allow a check to be written for ph, and therefore they must acquire a loan to finance this large expenditure.

In general, a mortgage loan requires a down payment equal to χ percent of the value of the house. The amount χph represents the amount of equity in the house at the time of purchase, and $D_0 = (1-\chi)ph$ represents the initial amount of the loan. In a particular period, denoted by n, the borrower faces a payment amount m_n (i.e., monthly or yearly payment) that depends on the size of the original loan, D_0, the length of the mortgage, N, and the mortgage interest rate, r^m. This payment can be subdivided into an amortization (or principal) component, A_n, which is determined by the amortization schedule, and an interest component, I_n, which depends on the payment schedule. That is,

(1) $$m_n = A_n + I_n, \quad \forall n,$$

where the interest payments are calculated by $I_n = r^m D_n$.[3] An expression that determines how the remaining debt, D_n, changes over time can be written as

(2) $$D_{n+1} = D_n - A_n, \quad \forall n.$$

This formula shows that the level of outstanding debt at the start of period n is reduced by the amount of any principal payment. A principal payment increases the level of equity in the home. If the amount of equity in a home at the start of period n is defined as H_n, a payment of principal equal to A_n increases equity in the house available in the next period to H_{n+1}. Formally,

(3) $$H_{n+1} = H_n + A_n, \quad \forall n,$$

where $H_0 = \chi ph$ denotes the home equity in the initial period.[4]

This representation of mortgage contracts is very general and summarizes many of the different contracts available in the financial markets. For example, this formulation can accommodate a no-down-payment loan by setting $\chi = 0$ so that the initial loan is equal to $D_0 = ph$. Because this framework can be used to characterize differences in the amortization terms and payment schedules, we use it to describe the characteristics of some prominent types of mortgage loans.

Mortgage Loans with Constant Payments

In the United States, fixed-rate mortgages (FRMs) typically are considered the standard mortgage contract. This loan product is characterized by a constant mortgage payment over the term of the mortgage, $m \equiv m_1 = \ldots = m_N$. This value, m, must be consistent with the condition that the present value of mortgage payments repays the initial loan. That is,

$$D_0 \equiv \chi ph = \frac{m}{1+r} + \ldots + \frac{m}{(1+r)^{N-1}} + \frac{m}{(1+r)^N}.$$

If this equation is solved for m, we can write

$$m = \lambda D_0,$$

where $\lambda = r^m[1-(1+r^m)^{-N}]^{-1}$. Because the mortgage payment is constant each period, and $m = A_t + I_t$, the outstanding debt decreases over time $D_0 > \ldots > D_n$. This means the fixed-payment contract front-loads interest rate payments,

$$D_{n+1} = \left(1+r^m\right)D_n - m, \quad \forall n,$$

and, thus, back-loads principal payments,

$$A_n = m - r^m D_n.$$

The equity in the house increases each period by the mortgage payment net of the interest payment component:

$$H_{n+1} = H_n + \left[m - r^m D_n\right], \quad \forall n.$$

We now present some examples to illustrate key properties of the FRM contract.

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[3] The calculation of the mortgage payment depends on the characteristics of the contract, but for all contracts the present value of the payments must be equal to the total amount borrowed,

$$D_0 \equiv \chi ph = \frac{m_1}{1+r} + \frac{m_2}{(1+r)^2} + \ldots + \frac{m_N}{(1+r)^N}.$$

[4] It is important to state that for the sake of simplicity this framework assumes no changes in house prices. If house prices are allowed to change, then the equity equation would have to allow for capital gains and losses.

Table 2

Characteristics of a Fixed-Rate Mortgage at 6 Percent*

Payment	Total payment ($)	Interest ($)	Principal ($)	Remaining principal ($)
1	1,178.74	973.51	205.23	199,794.77
2	1,178.74	972.51	206.23	199,588.54
120	1,178.74	812.98	365.76	166,655.59
181	1,178.74	686.89	491.85	140,625.26
219	1,178.74	587.23	591.51	120,049.79
240	1,178.74	523.73	655.01	106,940.84
251	1,178.74	487.89	690.95	99,521.83
360	1,178.74	5.71	1,173.03	0.00
Total	424,346.40	224,346.40	200,000.00	—

NOTE: *Based on 30-year maturity.

Example 1. Consider the purchase of a house with a total cost of ph = \$250,000 using a loan with a 20 percent down payment, $\chi = 0.20$; an interest rate of 6 percent annually; and a 30-year maturity. This mortgage loan is for \$200,000.[5] Table 2 illustrates the changes in interest and principal payments per month over the length of the mortgage contract.

The first two rows of Table 2 show the mortgage payment in the first and second months of the contract. The monthly payment on this mortgage is \$1,178.74. In the first period, \$973.51 of the monthly payment goes to interest rate payments. This means the principal payment is only \$205.23.[6] Now, let us consider the mortgage payment 10 years into the mortgage. Although the monthly payment does not change, the principal payment has increased to \$365.76 and the interest payment component has decreased to \$812.98. After 10 years, the homeowner has paid off only \$33,344.41 of the original \$200,000 loan. The month after the halfway point in the mortgage occurs at period 181. The interest payment component of the monthly payment still exceeds the principal payment. In payment period 219—18 years and 3 months into the contract—the principal component of the monthly payment finally exceeds the interest payment component. From this point forward, the principal payment will be larger than the interest payment. At the end of 20 years, or period 240, the principal component of the \$1,178.74 monthly payment is \$655.01. However, \$106,941.84 is still owed on the original \$200,000 loan. The outstanding loan balance does not drop below \$100,000 until payment period 251. With a standard 30-year mortgage contract, it takes nearly 22 years to pay off half the mortgage loan. The remaining half of the mortgage will be repaid in the final 8 years of this mortgage.

Example 2. Table 3 shows the standard 30-year mortgage contract if the mortgage interest rate increases from 6 percent to 7 percent. A 1 percent increase in the interest rate increases the monthly mortgage payment from \$1,178.74 to \$1,301.85—a \$123.11 increase. Furthermore, the increase in the interest rate results in additional back-loading of principal payments. After 10 years, less than \$30,000 of the original balance is paid off. The payment period when the prin-

5 Tables 2 through 9 apply to the following situation: house purchase price of \$250,000 with a down payment of 20 percent (total loan amount of \$200,000). Other parameters vary as noted in the individual examples.

6 This is the same example used in McDonald and Thornton (2008). The numbers presented here are slightly different because of a difference in interest rate calculation. McDonald and Thornton calculate the monthly interest rate as 0.06/12 = 0.005. We calculate the monthly interest as 1.06(1/12) –1 = 0.004868. This explains why our payments are slightly lower.

Table 3

Characteristics of a Fixed-Rate Mortgage at 7 Percent*

Payment	Total payment ($)	Interest ($)	Principal ($)	Remaining principal ($)
1	1,301.85	1,130.83	171.02	199,828.98
2	1,301.85	1,129.86	171.99	199,656.99
120	1,301.85	967.32	334.53	170,746.58
181	1,301.85	830.00	471.85	146,322.72
239	1,301.85	647.47	654.38	113,858.74
240	1,301.85	643.77	658.08	113,200.66
260	1,301.85	565.22	736.63	99,965.68
360	1,301.85	7.31	1,294.54	0.00
Total	468,666.00	268,666.00	200,000.00	—

NOTE: *Based on 30-year maturity.

cipal component exceeds the interest component does not occur until period 239. In fact, the outstanding balance will not drop below $100,000 until payment 260—9 months later than if the interest rate is 6 percent (as in Example 1).

This table clearly illustrates the impact of interest rate changes on a mortgage loan. If the total interest payments on the mortgage contract presented in Table 2 are compared with those in Table 3, the 1 percent increase in the interest rate results in $44,320 of additional mortgage payments over the life of the mortgage.

Mortgage with Constant Amortization

As seen in Tables 2 and 3, the FRM accrues little equity in the initial years of the mortgage because most of the mortgage payment services interest payments. Some buyers would benefit by a combination of an FRM and faster equity accrual. Can a mortgage contract be designed to allow accrual of more equity in the initial periods, and what properties would be involved in such a contract? A mortgage contract with this benefit is known as a constant amortization mortgage (CAM). This loan contract allows constant contributions toward equity in each constant amortization mortgage period; that is, the amortization schedule is $A_n = A_{n+1} = A$. Because the interest repayment schedule depends on the size of outstanding level of debt, D_n, and the loan term, N, the mortgage payment, m_n, is no longer constant over the duration of the loan. Formally, the constant amortization term is calculated by

$$A = \frac{D_0}{N} = \frac{(1-\chi)ph}{N}.$$

If the expression for the interest payments is used, the monthly mortgage payment, m_n, will decrease over the length of the mortgage. This characteristic of the CAM follows from the decline in outstanding principal over the life of the contract. The monthly payment is determined by

$$m_n = \frac{D_0}{N} + r^m D_n.$$

For this contract, the changes in the outstanding level of debt and home equity are represented by

$$D_{n+1} = D_n - \frac{D_0}{N}, \quad \forall n,$$

and

$$H_{n+1} = H_n + \frac{D_0}{N}, \quad \forall n.$$

Example 3. We consider a $250,000 30-year loan with a 20 percent down payment and a 6 percent annual interest rate to show the charac-

Table 4

Characteristics of a Constant Amortization Mortgage at 6 Percent*

Payment	Total payment ($)	Interest ($)	Principal ($)	Remaining principal ($)
1	1,529.07	973.51	555.56	199,444.44
2	1,526.36	970.81	555.56	198,888.89
120	1,207.27	651.71	555.56	133,333.33
156	1,109.92	554.36	555.56	113,333.33
181	1,042.31	486.76	555.56	99,444.44
240	882.76	327.21	555.56	66,666.67
360	558.26	2.70	555.56	0.00
Total	375,718.58	175,718.58	200,000.00	—

NOTE: *Based on 30-year maturity.

teristics of this type of contract. Table 4 presents the monthly mortgage payment, principal component, and interest component.

The monthly payment with this contract has a much different profile than that of a fixed-payment mortgage loan. Clearly, the amount of the mortgage payment declines over the life of the loan. The initial payment is nearly three times the size of the payment in the last period. Principal payments are constant over the life of the loan, thus allowing for faster equity accumulation. Half of the original principal is repaid halfway through the loan. From a wealth accumulation perspective, this is an attractive feature. However, the declining payment profile is not positively correlated with a normal household's earning pattern during the first half of the life cycle: Mortgage payments are highest when earnings tend to be lower. From a household budget perspective, this could be a very unattractive option.

Balloon and Interest-Only Loans

The key property of the CAM is the payment of principal every period. In contrast, balloon and interest-only loans allow no amortization of principal throughout the term of the mortgage. A balloon loan is a very simple contract in which the entire principal borrowed is paid in full in the last payment period, N. This product tends to be more popular when mortgage rates are high and home buyers anticipate lower future mortgage rates. In addition, homeowners who expect to stay in their homes only for a short time may find this contract attractive as they are not concerned about paying principal. The amortization schedule for this contract can be written as

$$A_n = \begin{cases} 0 & \forall n < N \\ (1-\chi)\,ph & n = N \end{cases}.$$

This means that the mortgage payment in all periods, except the last period, is equal to the interest rate payment, $I_n = r^m D_0$. Hence, the mortgage payment for this contract can be specified as

$$m_n = \begin{cases} I_n & \forall n < N \\ \left(1+r^m\right) D_0 & n = N \end{cases},$$

where $D_0 = (1-\chi)ph$. The evolution of the outstanding level of debt can be written as

$$D_{n+1} = \begin{cases} D_n, & \forall n < N \\ 0, & n = N \end{cases}.$$

With an interest-only loan and no change in house prices, the homeowner never accrues equity beyond the initial down payment. Hence, $A_n = 0$ and $m_n = I_n = r^m D_0$ for all n. In essence, the homeowner effectively rents the property from the lender and the mortgage (interest) payments are the effective rental cost. As a result, the monthly

Table 5

Characteristics of a Balloon Mortgage at 6 Percent*

Payment	Total payment ($)	Interest ($)	Principal ($)	Remaining principal ($)
1	973.51	973.51	0.00	200,000.00
2	973.51	973.51	0.00	200,000.00
180	973.51	973.51	0.00	200,000.00
181	1,670.59	973.51	697.08	199,302.92
219	1,670.59	832.25	838.34	170,141.84
240	1,670.59	742.26	928.33	151,562.86
290	1,670.59	487.16	1,183.43	98,898.87
360	1,670.59	8.09	1,662.50	0.00
Total	475,938.02	275,938.02	200,000.00	—

NOTE: *Based on 30-year maturity, 15 years interest only.

Table 6

Characteristics of an Adjustable-Rate Mortgage with a Constant Interest Rate of 6 Percent*

Payment	Total payment ($)	Interest ($)	Principal ($)	Remaining principal ($)
1	973.51	973.51	0.00	200,000.00
2	973.51	973.51	0.00	200,000.00
36	973.51	973.51	0.00	200,000.00
37	1228.20	973.51	254.89	199,745.30
120	1,228.20	847.10	381.10	173,648.03
181	1,228.20	715.71	512.49	146,525.31
219	1,228.20	611.86	616.34	125,086.37
240	1,228.20	545.70	682.50	111,427.30
257	1,228.20	486.97	741.23	99,303.08
360	1,228.20	5.95	1,222.25	0.00
Total	432,983.16	232,983.16	200,000.00	—

NOTE: *Based on 30-year maturity, 3 years interest only.

mortgage payment is minimized because no periodic payments toward equity are made. A homeowner is fully leveraged with the bank with this type of mortgage contract. If capital gains are realized, the return on the housing investment is maximized. If the homeowner itemizes tax deductions, a large interest deduction is an attractive by-product of this contract.

Example 4. This example illustrates a balloon contract with a 15-year interest-only loan that is rolled into a 15-year fixed-payment mortgage. Table 5 shows the payment profiles for this contract. We also assume an interest rate of 6 percent and a 20 percent down payment.

The interest-only part of the loan requires 180 mortgage payments of $973.51 just to cover

Table 7

Characteristics of an Adjustable-Rate Mortgage with a Rising Interest Rate*

Payment	Total payment ($)	Interest ($)	Principal ($)	Remaining principal ($)
1	973.51	973.51	0.00	200,000.00
2	973.51	973.51	0.00	200,000.00
36	973.51	973.51	0.00	200,000.00
37	1,347.72	1,130.83	216.89	199,783.11
120	1,347.72	1,001.40	346.32	176,762.45
181	1,347.72	859.24	488.48	151,477.91
239	1,347.72	670.28	677.44	117,869.91
240	1,347.72	666.45	681.27	117,188.65
264	1,347.72	567.74	779.98	99,630.97
360	1,347.72	7.57	1,340.15	0.00
Total	471,707.64	271,707.64	200,000.00	—

NOTE: *Based on 30-year maturity, 3 years interest only at a 6 percent interest rate, and the remaining years at 7 percent.

the interest obligations on the $200,000 loan. After 15 years, the mortgage payment increases to $1,670.59 because the 15-year balloon loan is rolled into a 15-year FRM. Payment number 219 denotes the month in which principal payments exceed interest payments. In period 290, half of the $200,000 debt will be paid off. With this type of mortgage contract, it takes more than 24 years to accrue $100,000 in equity.

Example 5. Some ARMs used in recent years have a very short period of interest-only payments. Table 6 presents the payment profiles for a 3-year interest-only ARM that rolls into a 27-year standard FRM. The assumptions for the interest rate, total contract length, and down payment remain unchanged.

The monthly interest payments for this interest-only ARM are $973.51. Once the standard 27-year contract takes effect, the monthly mortgage payment increases by $254.69 to $1,228.20. This increase is not caused by an interest rate increase, but rather payment toward principal.

Example 6. Mortgage interest rates have begun to increase recently. What effect does this have on an interest-only ARM? To show this effect, we allow the interest rate to increase to 7 percent for the standard FRM that is obtained after the 3-year ARM expires. Table 7 presents the various payment patterns. A 100-basis-point increase in the interest rate causes the monthly payment to increase to $1,347.72 from $1,228.20—a 38 percent increase in the mortgage payment from the interest-only payments. This example illustrates the risk facing homeowners when the interest rate increases before the transition to a standard FRM.

Graduated-Payment Mortgages

The repayment structures of the previous contract examples are relatively rigid. Payments are either constant during the entire contract or proportional to the outstanding level of debt. Mortgage contracts can be designed with a variable repayment schedule. This section focuses on mortgage loan payments that increase over time, $m_1 < \ldots < m_N$. This feature could attract first-time buyers because payments are initially lower than payments in a standard contract. When a buyer's income grows over the life cycle, this loan product allows for stable housing expenditure as a ratio to income. However, the buyer's equity in the home builds at a slower rate than with the standard contract, which may explain this product's lack of popularity historically. Mortgage contracts with

Table 8

Characteristics of a Graduated-Payment Mortgage: 1 Percent Geometric Growth Rate*

Payment	Total payment ($)	Interest ($)	Principal ($)	Remaining principal ($)
1	195.18	973.51	–778.33	200,778.33
2	197.13	977.30	–780.17	201,558.50
120	637.79	1,459.98	–822.19	300,763.84
181	1,170.26	1,666.83	–496.57	342,933.91
220	1,725.11	1,719.49	5.57	353,260.70
240	2,104.96	1,701.52	403.44	349,161.20
344	5,924.70	508.34	5,416.36	99,017.59
360	6,947.18	33.65	6,913.53	0.00
Total	682,149.10	482,149.10	200,000.00	—

NOTE: *Based on interest rate of 6 percent, 30-year maturity, and a payment growth of 1 percent.

variable repayment schedules are known as graduated-payment mortgages (GPMs). These contracts are of special interest because their features are similar to those of mortgage contracts sold to subprime borrowers.

The repayment schedule for a GPM depends on the growth rate of these payments. The growth rate of payments is specified in the mortgage contract, and borrowers considering this contract must know this condition. We present examples to illustrate why knowledge of this parameter or condition is important. Typical GPM growth patterns are either geometric or arithmetic. We focus on GPMs with geometric growth patterns.

With this type of contract, mortgage payments evolve according to a constant geometric growth rate denoted by

$$m_{n+1} = (1+g)m_n,$$

where $g > 0$. This means the amortization and interest payments also increase as

$$m_n = A_n + I_n.$$

The initial mortgage payment is determined by

$$m_0 = \lambda_g D_0,$$

where $\lambda_g = (r^m - g)[1 - (1 + r^m)^{-N}]^{-1}$. The law of motion for the outstanding debt satisfies

$$D_{n+1} = (1+r^m)D_n - (1+g)^n m_0,$$

and the amortization term is $A_n = \lambda_g D_0 - r^m D_n$.

Example 7. Table 8 shows the implications for payments of a GPM contract when the mortgage payments grow at 1 percent per payment. We maintain the assumption of a 30-year contract with a 20 percent down payment and a 6 percent annual interest rate.

Clearly, the initial payments of this mortgage are very low, which explains why this contract is attractive for first-time buyers. However, these low payments come at a cost: The monthly payment does not cover the interest on the outstanding balance. Thus, the remaining principal increases. This mortgage contract exhibits negative amortization. In this example, the mortgage payment does not cover the interest on the principal for the first 219 months. The maximum remaining principal for this home purchase increases to more than $350,000 from the original $200,000 debt. It is interesting to note that the final $100,000 principal is paid in the final 16 months of this mortgage. Because the principal is back-loaded and must be paid off, the monthly payment must increase over time. The monthly mortgage payment tops out in the last month of the contract at $6,913.53. A homeowner who chooses this con-

Table 9

Characteristics of a Graduated-Payment Mortgage: 0.1 Percent Geometric Growth Rate*

Payment	Total payment ($)	Interest ($)	Principal ($)	Remaining principal ($)
1	1,030.68	973.51	57.17	199,942.83
2	1,031.71	973.23	58.48	199,884.36
120	1,160.85	884.19	276.67	181,372.92
240	1,308.78	614.19	694.59	125,485.59
273	1,352.67	489.90	862.77	99,782.99
360	1,475.56	7.15	1,468.41	0.00
Total	446,356.77	246,356.77	200,000.00	—

NOTE: *Based on interest rate of 6 percent, 30-year maturity, and a payment growth of 0.1 percent.

tract pays $482,149.10 in total interest payments. Compared with the FRM contract presented in Table 2, total interest payments are more than double. These characteristics make GPMs risky from a lender's perspective because the potential for default is greater, which is one reason this type of contract has not historically been a factor in the mortgage market.

Example 8. Table 9 shows the importance of the payment growth parameter by reducing the monthly growth rate from 1 percent to 0.1 percent. Negative amortization does not occur with a lower monthly growth rate. Perhaps the most striking result is the amount of total interest payments over the length of the mortgage contract. When the mortgage contract has a 1 percent monthly growth rate, total interest payments are $482,149.10. If the monthly growth rate falls to 0.1 percent, total interest payments are $246,356.77. Clearly there is a cost to loans with negative amortization.

Combo Loans

In the late 1990s a new mortgage product became popular as a way to avoid large down payments and mortgage insurance.[7] This product is known as the combo loan and amounts to two different loans. Different types of CLs are offered in the mortgage industry; for example, an 80-15-5 loan implies a primary loan for 80 percent of the value, a secondary loan for 15 percent, and a 5 percent down payment. Another example is the so-called no-down-payment, or an 80-20 loan, which consists of a primary loan with a loan-to-value ratio of 80 percent and a second loan for the 20 percent down payment.

Formally, the primary loan covers a fraction of the total purchase, $D_1 = (1-\chi)ph$, with a payment schedule, m_n^1, and maturity, N_1. The second loan partially or fully covers the down payment, $D_2 = \vartheta\chi ph$, where $\vartheta \in (0,1]$ and represents the fraction of the down payment financed by the second loan. The second loan has an interest premium $r_2^m = r_1^m + \zeta$ (where $\zeta > 0$), a mortgage payment m_n^2, and a maturity $N_2 \le N_1$. In this case,

$$m_n = \begin{cases} m^1 + m^2 & \text{when } N_2 \le n \le N_1 \\ m^1 & \text{when } n < N_2 \end{cases}.$$

Because both loans are of a fixed-rate form, the laws of motion are equivalent to those presented in the FRM contract discussion. Table 10 shows characteristics of a CL.

Example 9. Table 10 presents the profile for an 80-20 CL for our $250,000 house. The first $200,000 is borrowed with the standard fixed-payment mortgage at 6 percent interest. The remaining $50,000 is financed using another

[7] Government-sponsored enterprises (GSEs) initiated the use of this product in the late 1990s and it became popular in private mortgage markets between 2001 and 2002.

Table 10

Characteristics of an 80-20 Combo Loan Mortgage*

Payment	Total payment ($)	Interest ($)	Principal ($)	Remaining principal ($)
1	1,535.94	1,295.21	240.73	249,759.27
2	1,535.94	1,293.98	241.96	249,517.32
120	1,535.94	1,094.04	441.90	210,261.45
181	1,535.94	931.49	604.45	178,528.13
156	1,535.94	554.36	555.56	113,333.33
228	1,535.94	765.78	770.16	146,301.19
240	1,535.94	716.53	819.41	136,742.23
281	1,535.94	522.76	1,013.15	99,220.31
360	1,535.94	7.99	1,527.95	0.00
Total	552,938.40	302,938.40	250,000.00	—

NOTE: *Based on interest rate of 6 percent, 30-year maturity, and second loan rate of 8 percent.

fixed-payment mortgage that incorporates a risk premium of 2 percent. We will assume the second mortgage is also for 30 years. (In reality, the second mortgage is usually for 10 or 15 years.) The second loan for $50,000 increases the monthly payment by $357.20. The mortgage payment pattern of this CL is very similar to the basic fixed-payment mortgage. This is not surprising because the CL is nothing more than a combination of two FRMs. An obvious question for borrowers is why they should not obtain just one FRM with no down payment. The larger single loan would require mortgage insurance. The total monthly payment, including the mortgage insurance, would exceed the monthly payment on the CL. The CL is attractive for one segment of buyers who desire to enter the housing market: young buyers with high incomes. These buyers can afford the mortgage payment, but they have not yet had time to accumulate savings for the down payment.

A MODEL OF HOUSING DECISIONS AND MORTGAGE CHOICES

The previous section described various features and properties of mortgage contracts available in the marketplace. However, the discussion did not detail the characteristics of individuals who might choose a particular contract. In addition, no mention was made of the ramifications of alternative contracts for the performance of the aggregate economy. The only way to discuss these issues is by analyzing alternative mortgages in the context of a model economy in which buyers can choose from among a set of mortgage products. In this section, we use a simplified version of the consumer problem used by Chambers, Garriga, and Schlagenhauf (2007a,b) to address the implications of mortgage choice for the performance of the aggregate economy (i.e., house prices, interest rates). This model allows us to focus on how types of mortgages influence the homeownership decision. This modeling style allows quick analysis of aggregate implications of mortgage markets and yet maintains the details needed to identify implications across different income, wealth distribution, and age cohorts.

Model Features

Age Structure. We develop a life cycle model with ex ante heterogeneous individuals. Let j denote the age of an individual and let J represent the maximum number of periods an individual can live. At every period, an individual faces mortality risk and uninsurable labor-earning uncertainty. The survival probability, conditional on the individual being alive at age j, is denoted by $\psi_{j+1} \in [0,1]$, with $\psi_1 = 1$ and $\psi_{j+1} = 0$. Earning uncertainty implies that the individual is subject to income shocks that cannot be insured via private contracts. In addition, we assume that annuity markets for mortality risk are absent. The lack of these insurance markets creates a demand for precautionary savings to minimize fluctuations in consumption goods, c, and in the consumption of housing services, s, over the life cycle.

Preferences. Individual preferences rank goods (consumption and housing) according to a utility function, $u(c,s)$. The utility function has the property that additional consumption increases utility and also results in declining marginal utility. Consumption over periods is discounted at rate β and, thus, lifetime utility is defined as

$$v_1 = E\sum_{j=1}^{J}\psi_j\beta^{j-1}u\left(c_j,s_j\right).$$

The assumption that utility is separable over time allows for a simple recursive structure of preferences for every realization of uncertainty:

$$v_1 = u\left(c_1,s_1\right)+\beta E\sum\nolimits_{j=2}^{J}\psi_j\beta^{j-2}u\left(c_j,s_j\right).$$

Using the definition of expected lifetime utility, we can write the previous expression as

$$v_1 = u\left(c_1,s_1\right)+\beta E v_2,$$

where

$$v_2 = \sum\nolimits_{j=2}^{J}\psi_j\beta^{j-2}u\left(c_j,s_j\right)$$

represents the future lifetime expected utility.

Asset Structure. Individuals have access to a portfolio of assets to mitigate income and mortality risk. We consider two distinct assets: a riskless financial asset denoted by a' with a net return r and a risky housing durable good denoted by h' with a market price, p, where the prime is used to denote future variables. To keep things simple, we assume that the price of housing does not change over time, so $p = p'$. This assumption simplifies the problem because households do not need to anticipate changes in house prices. A housing investment of size h' can be thought of as the number of square feet in the house. A house of size h' yields s services.[8] If a household does not invest in housing, $h = 0$, the household is a renter and must purchase housing services from a rental market. The rental price of a unit of housing services is R.

Housing investment is financed through long-term mortgage contracts and is subject to transaction costs. We need to summarize the information required so that the monthly payment, remaining principal, and equity position in the house can be identified for any mortgage contract. This critical information consists of the house size, h, the type of mortgage contract, z, and the remaining length of the mortgage, n. This information set can be used to identify the desired information concerning a household's mortgage contract.

Household Income. Household income varies over the buyer's life cycle and depends on whether the individual is a worker or a retiree. For workers younger than the mandatory retirement age, $j < j^*$, income is stochastic and depends on the basic wage income, w, a life cycle term that depends on age, υ_j, and the persistent idiosyncratic component, ε, drawn from a probability distribution that evolves according to the transition law, $\Pi_{\varepsilon,\varepsilon'}$. For an individual older than j^*, a retirement benefit, θ, is received from the government equal to θ. Income can be specified as

$$y\left(a,\varepsilon,j,\upsilon_j\right) = \begin{cases} w\varepsilon\upsilon_j+\left(1+r\right)a, & \text{if} \quad j<j^*, \\ \theta+\left(1+r\right)a, & \text{if} \quad j\geq j^*. \end{cases}$$

In the presence of mortality risk and missing annuity markets, we assume borrowing constraints, $a' \geq 0$, to prevent individuals (buyers and renters) from dying with negative wealth. We also assume that households are born with initial wealth dependent on their initial income level.

The Decision Problem. Individuals make decisions about consumption goods, c, housing services, s, a mortgage contract type, z, and

8 For the sake of simplicity, we assume a linear relationship between house size and services generated. In other words, $s = h'$.

$$\textbf{Current renter: } h=0 \quad \begin{bmatrix} \text{Continues renting: } h'=0 \\ \text{Purchases a house: } h'>0 \end{bmatrix}$$

$$\textbf{Current owner: } h>0 \quad \begin{bmatrix} \text{Stays in house: } h'=h \\ \text{Changes size (upsize or downsize): } h'\neq h \\ \text{Sell and rent: } h'=0 \end{bmatrix}$$

investment in assets, a', and housing, h'. The individual's current-period budget constraint depends on the household's asset holdings, the current housing investment, the remaining length of the mortgage, labor income shock, and age. We can isolate five possible decision problems that a household must solve.

The household value function, v, is described by a vector of so-called state variables that provide sufficient information of the position of the individual at the start of the period. The state vector is characterized by the level of assets at the start of the period, a, the prior-period housing position, h, the number of periods remaining on an existing mortgage, n, the mortgage contract type, z, the value of the current-period idiosyncratic shock, ε, and the age of the individual, j. To shorten notation of the individual's characteristics, we define $x = (a,h,n,z,\varepsilon,j)$. Using a recursive approach, we know that the household decisions for c,s,z,a' and h' depend on the x vector. For example, suppose that x contains the following information, $x = (1000,2000,56,FRM,2,36)$. This vector tells us that the individual has \$1,000 of non-housing wealth, a 2,000-square-foot home with a market value given by $p \times 2{,}000$, where p represents the given price per square foot, 56 pending mortgage payments with the bank, an FRM, the income shock this period is two times average income, and the individual's age is 36. The decisions made by this individual will differ from those of an individual who has a different state vector $x = (20000,2000,56,FRM,2,41)$, because the second individual has more wealth and is 5 years older. For individuals who do not own a home, the information vector would have many zero entries, such as $x = (a,0,0,0,\varepsilon,j)$.

Given all the possible options, the individual could be in one of five situations with respect to the housing investment and mortgage choice decisions. These five decisions are summarized in the box above.

We now detail the various decision problems. First, we consider an individual who starts as a renter and then consider an individual who starts as a homeowner.

Renters. An individual who is currently renting ($h = 0$) has two options: continue renting ($h' = 0$) or purchase a house ($h' > 0$). This is a discrete choice in ownership that can easily be captured by the value function, v (present and future utility), associated with these two options. Given the relevant information vector $x = (a,0,0,0,\varepsilon,j)$, the individual chooses the option with the higher value, which can be expressed as

$$v(x) = \max\left\{v^r, v^o\right\}.$$

The value associated with continued renting is determined by the choice of consumption goods, c, housing services, s, and investment in assets, a', which solves the problem

$$\begin{aligned} v^r(x) &= \max u(c,s) + \beta_{j+1} E v(x'), \\ \text{s.t. } & c + a' + Rs = y(x). \end{aligned}$$

The decisions are restricted to positive values for c,s,a' and the evolution of the state vector that summarizes the future information as given by $x' = (a',0,0,0,\varepsilon',j+1)$, where a' denotes next period's wealth, ε' represents next period's realization of the income shock, and $j+1$ captures the fact that the individual will be one period older.

The renter who chooses to purchase a house must solve a different problem as choices must now be made over $h' > 0$, a mortgage type, z, as well as c, s, and a'. This decision problem can be written as

$$v^o(x) = \max u(c,s) + \beta_{j+1} Ev(x'),$$
$$\text{s.t. } c + a' + [\varphi_b + \chi(z')]ph' + m(x) = y(x).$$

It should be noted that a purchase of a house requires two up-front expenditures: transaction costs (i.e., realtors' fees, closing costs) that are proportional to the value of the house, $\varphi_b ph'$, and a down payment to the mortgage bank for a fraction of the value of the house, $\chi(z')$ (i.e., 20 percent of the purchase price). These payments are incurred only at the time of the purchase. Homeowners also must make mortgage payments. These payments are denoted by $m(x)$ and depend on relevant variables, such as the loan amount, $(1-\chi)ph'$, the type of mortgage (i.e., FRM vs. ARM), the length of the contract (i.e., 30 or 15 years), and the interest rate associated with the loan. It is important to restate that a homeowner who purchases a house of size h' receives s units of housing consumption. The value of these housing services is denoted by Rs^h. This value does not appear in the budget constraint because these services are consumed internally. As a result, the value of services generated is canceled by the value of services consumed internally. The household's decisions influence the information state in the following period; that is, $x' = (a',h',N(z)-1,z',\varepsilon',j+1)$. Again, to determine whether an individual continues to rent or purchases a home, we need to solve both problems—$v^r(x)$ and $v^0(x)$— and find the one that yields the highest value. When $v^r(x) > v^0(x)$, the individual continues to rent; otherwise he or she will become a homeowner.

Owners. The decision problem for an individual who currently owns a house, $(h > 0)$, has a similar structure. However, a homeowner faces a different set of options: stay in the same house, $(h' = h)$, purchase a different house, $(h' \neq h)$, or sell the house and acquire housing services through the rental market, $(h' = 0)$. Given the relevant information $x = (a,h,n,z,\varepsilon,j)$, the individual solves

$$v(x) = \max\{v^s, v^c, v^r\}.$$

Each of these three different values is calculated by solving three different decision problems. If the homeowner decides to stay in the current house, the optimization problem can be written as

$$v^s(x) = \max u(c,h') + \beta_{j+1} Ev(x')$$
$$\text{s.t. } c + a' = y(x) - m(x).$$

This problem is very simple, because the homeowner must make decisions only on consumption and saving after making the mortgage payment. If the schedule of pending mortgage payments shows zero, $n = 0$, then the implied mortgage payment is also set to zero, $m(x) = 0$. The future state of information for this case is given by $x' = (a',h',n', z',\varepsilon',j+1)$, where $n' = \max\{n-1,0\}$.

For the homeowner who decides to either upsize or downsize, $(h \neq h')$, the problem becomes

$$v^c(x) = \max u(c,h') + \beta_{j+1} Ev(x')$$
$$\text{s.t. } c + a' + [\varphi_b + \chi(z')]ph' + m(x)$$
$$= y(x) + [(1-\varphi_s)ph - D(n,z)].$$

This individual must sell the existing property to purchase a new one. The choices depend on the income received from selling the property, ph, net of transactions costs from selling, φ_s, and the remaining principal, $D(n,z)$, owed to the lender. The standing balance depends on whether the mortgage has been paid off ($n = 0$ and $D(n,z) = 0$) or not ($n > 0$ and $D(n,z) > 0$) and the type of loan contract. For example, mortgage loans with a slow amortization usually imply large remaining principal when the property is sold over the length of the loan, whereas contracts such as the constant amortization imply a much faster repayment. A new loan, z', must be acquired if the individual upsizes and purchases a new house, $h' > 0$. The relevant future information is given by $x' = (a',h',N-1, z',\varepsilon',j+1)$.

Finally, we solve the problem of a homeowner who sells the house, $h > 0$, and becomes a renter, $h' = 0$.[9] The optimization problem is very similar to the previous one. However, in this case the individual must sell the home and rent, Rs. Formally,

[9] In the last period, all households must sell h, rent housing services, and consume all their assets, a, as a bequest motive which is not in the model. In the last period, $h' = a' = 0$.

$$v^r(x) = \max u(c,s) + \beta_{j+1} E v(x'),$$
$$\text{s.t. } c + a' + Rs = y(x) + \left[(1-\varphi_s)ph - D(n,z)\right],$$

where the relevant future information is simply given by $x' = (a',0,0,0,\varepsilon',j+1)$.

Given the initial information summarized in x, the choice of whether to stay in the house, change the housing size, or sell the house and become a renter depends on the values of v^s, v^c, and v^r.

Aggregation and Parameterization

We want our model economy to be consistent with certain features of the U.S. economy. In particular, we want to ensure that the choices of functional forms and parameter values are consistent with key features of the U.S. housing market. Replicating these features requires aggregating the microeconomic behavior of all the households in the economy. Because individuals are heterogeneous along six different dimensions—level of wealth, housing holdings, pending mortgage payments, type of mortgage used to finance the house, income shock, and age—our aggregation needs to take into account the number of individuals who have the same characteristics and the sum across these characteristics. To aggregate these dimensions, we define $\Phi(x)$ as the fraction of individuals who have a given level of characteristics $x = (a,h,n,z,\varepsilon,j)$.

We can calculate aggregate statistics of the economy by taking the weighted average of all the household-level decisions across characteristics. As an example we would generate the aggregate housing stock, wealth, and consumption of housing services (or square feet) by calculating

$$H = \int h'(x)\Phi(dx);$$
$$W = \int a'(x)\Phi(dx);$$
$$S = \int s(x)\Phi(dx).$$

The model can generate other aggregates of interest in a similar manner. We start by discussing how the model is parameterized.

Demographics. A period in the model is taken to be three years. Individuals enter the labor force at age 20 (model period 1) and potentially live until age 86 (model period 23). Retirement is assumed to be mandatory at age 65 (model period 16). Individuals survive to the next period with probability ψ_{j+1}.[10] The size of the age-specific cohorts, μ_j, needs to be specified. Because of our focus on steady-state equilibrium, these shares must be consistent with the stationary population distribution. As a result, these shares are determined from $\mu_j = \psi_j \mu_{j-1}/(1+\rho)$ for $j = 1,2,\ldots,J$ and

$$\sum_{j=1}^{J} \mu_j = 1,$$

where ρ denotes the population growth rate. Using the resident population as the measure of the population, the annual growth rate is set at 1.2 percent.

Functional Forms. The choice of preferences is based on empirical evidence. The first-order condition that determines the optimal amount of housing consumption is denoted by

$$\frac{u_{s_j}}{u_{c_j}} = R,$$

where at the optimum $s_j = h'_j$. Jeske (2005) documents that the h_j/c_j ratio is increasing by age j. He points out that standard homothetic preferences over consumption and housing services,

$$u(c_j,s_j) = \left[\gamma c_j^{\sigma} + (1-\gamma)s_j^{\sigma}\right]^{\frac{1}{\sigma}},$$

imply a constant ratio

$$\frac{h_j}{c_j} = \left(\frac{(1-\gamma)}{\gamma R}\right)^{\frac{1}{1-\sigma}},$$

because the parameters γ and σ and the rental price R do not vary across age. Therefore, this preference specification is inconsistent with the empirical evidence over the life cycle. A preference structure consistent with the evidence is denoted by

$$u(c,s) = \gamma\frac{c^{1-\sigma_1}}{1-\sigma_1} + (1-\gamma)\frac{s^{1-\sigma_2}}{1-\sigma_2},$$

where the implied first-order condition is denoted by

[10] These probabilities are set at survival rates observed in 1994, and the data are from the National Center for Health Statistics (U.S. Department of Health and Human Services, 1994).

$$\frac{h_j^{\sigma_2}}{c_j^{\sigma_1}} = \frac{(1-\gamma)}{\gamma R}.$$

This expression represents a nonlinear relationship between h_j and c_j that varies by age, j. The coefficients, σ_1 and σ_2, determine the curvature of the utility function with respect to consumption and housing services. The relative ratio of σ_1 and σ_2 determines the growth rate of the housing-to-consumption ratio. A larger curvature in consumption relative to the curvature in housing services implies that the marginal utility of consumption exhibits relatively faster diminishing returns. When household income increases over the life cycle (or different idiosyncratic labor income shocks), a larger fraction of resources is allocated to housing services. We set $\sigma_2 = 1$ and $\sigma_2 = 3$ to match the observed average growth rate while the preference parameter γ is estimated. The discount factor, β, is set at 0.976, which is derived from Chambers, Garriga, and Schlagenhauf (2007a).

Endowments. Workers are assumed to have an inelastic labor supply, but the effective quality of their supplied labor depends on two components. One component is age specific, υ_j, and is designed to capture the hump in life cycle earnings. We use U.S. Census Bureau (1994) data to construct this variable. The other component captures the stochastic component of earnings and is based on Storesletten, Telmer, and Yaron (2004). We discretize this income process into a five-state Markov chain using the methodology presented by Tauchen (1986). Our reported values reflect the three-year horizon used in the model. As a result, the efficiency values associated with each possible productivity value, ε, are

$$\varepsilon \in \mathbf{E} = \{4.41, 3.51, 2.88, 2.37, 1.89\},$$

and the transition matrix is

$$\pi = \begin{bmatrix} 0.47 & 0.33 & 0.14 & 0.05 & 0.01 \\ 0.29 & 0.33 & 0.23 & 0.11 & 0.03 \\ 0.12 & 0.23 & 0.29 & 0.24 & 0.12 \\ 0.03 & 0.11 & 0.23 & 0.33 & 0.29 \\ 0.01 & 0.05 & 0.14 & 0.33 & 0.47 \end{bmatrix}.$$

Each household is born with an initial asset position. This assumption accounts for the fact that some of the youngest buyers who purchase housing have some wealth. Failure to allow for this initial asset distribution creates a bias against the purchase of homes in the earliest age cohorts. As a result, we use the asset distribution observed in *Panel Study on Income Dynamics* (Institute for Social Research, 1994) to match the initial distribution of wealth for the cohort of age 20 to 23. Each income state has assigned the corresponding level of assets to match the nonhousing wealth-to-earnings ratio. We choose the basic level of earnings, w, as a scaler to match labor earnings over total earnings.

Housing. The housing market introduces a number of parameters. The purchase of a house requires a mortgage and down payment. In this paper, we focus on the 30-year FRM as the benchmark mortgage. As a result of the assumption that a period is three years, we set the mortgage length, N, to 10 periods. The down payment, χ, is set to 20 percent (matching facts from the 2004 U.S. Department of Commerce *American Housing Survey*, AHS). Buying and selling property is subject to transaction costs. We assume that all these costs are paid by the buyer and set $\sigma_s = 0$ and $\sigma_b = 0.06$.

Because of the lumpy nature of the housing investment (i.e., movement from $H = 0$ to $H > 0$), the specification of the second point in the housing grid has important ramifications. This grid point, $\underline{h}$, determines the minimum house size and has implications for the timing of investments in housing, wealth portfolio decisions, and the homeownership rate. We determine the size of this grid point as part of the estimation problem to avoid inadvertent implications on the results caused by this variable.

Estimation

We estimate five parameters using an exactly identified method of moments approach. The parameters that need to be estimated are the interest rate, r, the rental rate for housing, R, the price of housing, p, the wage rate w, and the size of the smallest housing investment position, $\underline{h}$. We identify these parameter values so that the

Table 11

Method of Moments Estimates*

Statistic	Data	Model estimate	Percent error
1. Ratio of wealth to gross domestic product	2.541	2.549	0.314
2. Ratio of housing stock to fixed capital stock	0.430	0.4298	–0.047
3. Ratio housing services to consumption of goods	0.230	0.235	2.7
4. Labor earnings over total earnings	0.700	0.71	1.4
5. Homeownership rate	0.640	0.643	0.468

Parameter	Value
1. Interest rate, r	0.0546
2. Rental price, R	0.3403
3. Housing price, p	1.4950
4. Wage rate, w	0.8768
5. Minimum house size, $\underline{h}$	1.4480

NOTE: *Values in annual terms.

resulting aggregate statistics in the model economy are equal to five targets observed in the U.S. economy.

i. **Wealth-to-gross income ratio (W/I)**. We find the target is the ratio of nonhousing wealth to gross income, which is about 2.541 (annualized value), for the period 1958-2001.

ii. **Housing stock-to-wealth ratio (H/W)**. For this ratio, the housing capital stock is defined as the value of fixed assets in owner and tenant residential property. The housing stock data are from the fixed asset tables of the Bureau of Economic Analysis (1958-2001) The ratio of the housing stock to nonhousing wealth is 0.43.

iii. **Housing services-to-consumption of goods ratio (RS/C)**. The targeted housing consumption-to-nonhousing consumption ratio is also based on *National Income and Product Accounts* (NIPA) data (1958-2001), where housing services are defined as personal consumption expenditure for housing while nonhousing consumption is defined as nondurable and services consumption expenditures net of housing expenditures (U.S. Department of Commerce, NIPA tables). The targeted ratio for 1994 is 0.23, but the number does not vary greatly over the period 1990-2000. This value is from Jeske (2005).

iv. **Labor earnings over total earnings**. The evidence from NIPA suggests that labor share of the economy is about 70 percent. We determine the value of w to match this observation.

v. **Homeownership ratio**. This target is based on data from the AHS (1994) for 1994 and is equal to 64.0 percent.

Table 11 summarizes the parameter estimates and the empirical targets. The moments and the parameter values are presented in annual terms. The model nicely matches the moments of the U.S. economy.

Model Evaluation

We can now take a more in-depth look at the results from a distribution perspective. We begin by studying the homeownership rate across both the age and the income distribution (Table 12).

Another dimension of interest is the consumption of housing services. We measure average

Table 12

Homeownership Rates by Age

	Homeownership rate (percent)					
Variable	Total	20-34 years	35-49 years	50-64 years	65-74 years	75-89 years
Data 1994	64.0	40.0	64.5	75.2	79.3	77.4
Baseline model 1994	64.3	37.1	80.6	81.5	81.5	62.5

SOURCE: U.S. Census Bureau, *Housing Vacancies and Homeownership* (1994) and U.S. Department of Commerce, *American Housing Survey* (1994).

Table 13

Owner-Occupied Housing Consumption by Age

	House size (square footage)					
Simulation	Total	20-34 years	35-49 years	50-64 years	65-74 years	75-89 years
Data 1994	2,137	1,854	2,220	2,301	2,088	2,045
Baseline model 1994	1,896	2,013	1,787	1,736	2,242	2,452

SOURCE: U.S. Department of Commerce, *American Housing Survey* (1994).

consumption of housing services by computing the average size of an owner-occupied house. Data from the AHS indicate the average owner-occupied house is 2,137 square feet. Our model implies an average house size of 1,895 square feet. Table 13 shows observed housing size by age cohorts. The model reasonably estimates homeowners' acquisition of appropriately sized homes. The average size for most age cohorts is within a few hundred square feet. Home size increases with age of the homeowner, which is observed only until age 65 in the data.[11]

Because households make savings decisions with respect to assets, the portfolio allocations implied by the model can be analyzed. In the model, a household's financial portfolio consists of asset holding and equity in housing investment. We use data from the 1994 *Survey of Consumer Finances* (Board of Governors, 1998) to determine the importance of housing in household portfolios. We define assets as bond and stock holdings, and housing is defined as the estimated value of an existing homeowner's house adjusted for the remaining principal. The data indicate housing represents a large fraction of a household's portfolio in the youngest age cohorts. This fraction declines with household age until around retirement age and then increases as retirees consume their nonhousing wealth after retirement. As shown in Figure 1, our model generates a similar pattern.

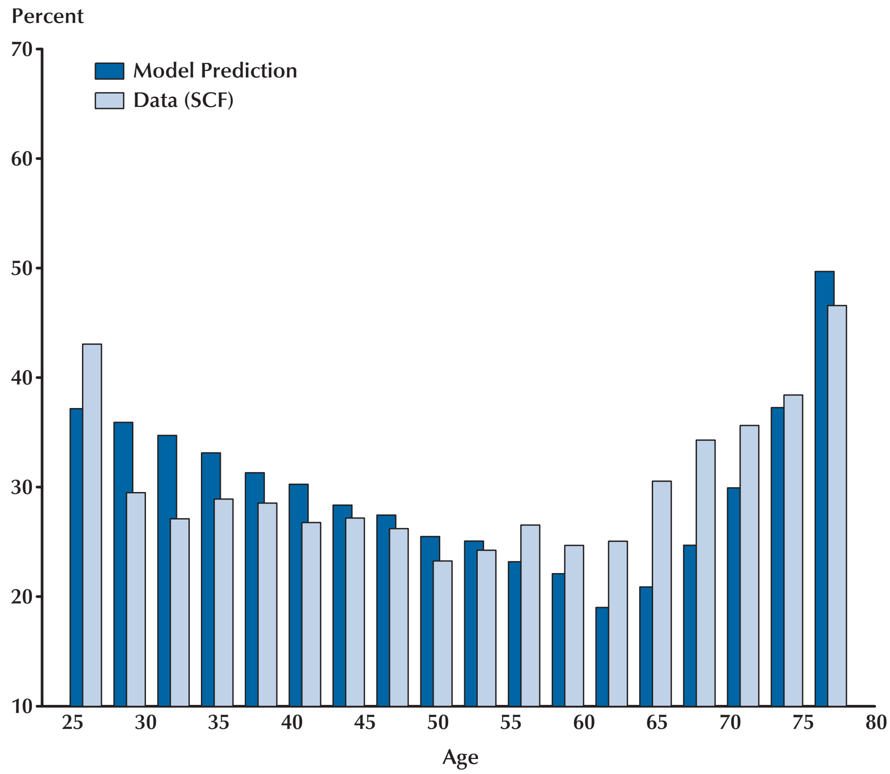
MORTGAGE CHOICES

In this section we look at the implications of mortgage innovation on the housing market, especially with regard to the rate of homeownership. We focus on two of the largest mortgage innovations: ARM-type and CL mortgage contracts. In the first example, households face an additional decision regarding the type of mortgage to finance their home purchase. We will allow a potential home buyer to choose between a 30-year fixed-

[11] It should be noted that the full equilibrium model with landlords in Chambers, Garriga, and Schlagenhauf (2007a,b) does capture the hump-shaped pattern in home size.

Figure 1

Housing in the Portfolio by Age



SOURCE: *Survey of Consumer Finances* (Board of Governors of the Federal Reserve System, 1998).

payment mortgage with a 20 percent down payment and an ARM with 3 years of interest-only payments followed by a 27-year fixed-payment mortgage. This simulation generates an aggregate homeownership rate of 65.83 percent, which is an increase of 1.5 percent from the baseline simulation. The effects are even more dramatic for homeownership rates by age.

Table 14 shows a very similar pattern to the baseline case with a few important differences. The biggest difference is the large increase in homeownership by the youngest cohort. For households younger than age 35, homeownership has surged to nearly 50 percent. Some of this increase in ownership is offset by a slight decrease in ownership later in life. This difference is explained by the labor income shocks for some of those who became owners by using ARM mortgages. The decision to own a home early in life delays the accumulation of capital assets, which insures the homeowner against bad income shocks. The average house size in this economy is 1,759 square feet. This implies that the introduction of ARMs leads to a large increase in the purchase of smaller homes, which tend to be purchased by lower-income households, who tend to be more exposed to labor income shocks. Without protection against income shocks, some homeowners are unable to make mortgage payments and thus become renters.

When considering mortgage finance and selection, we find that 51.7 percent of homeowners have some form of mortgage debt. As for the type of mortgage, 35.5 percent have a fixed-pay-

Table 14

Homeownership (Including an ARM) by Age

	Homeownership rate (percent)					
Simulation	Total	20-34 years	35-49 years	50-64 years	65-74 years	75-89 years
Benchmark	64.3	37.1	80.6	75.2	79.3	77.4
Model	65.8	49.1	80.3	76.3	72.9	64.7

SOURCE: Data generated by the model.

Table 15

Homeownership (Including Combo Loans) by Age

	Homeownership rate (percent)					
Simulation	Total	20-34 years	35-49 years	50-64 years	65-74 years	75-89 years
Benchmark model 1994	64.3	37.1	80.6	81.5	81.5	62.5
Model	68.6	42.2	88.0	81.6	83.2	66.9

SOURCE: Data generated by the model.

ment mortgage and 16.2 percent have an ARM mortgage. Only households with ARMs are in the bottom quintile of income distribution. ARMs can be attractive to many homeowners, but those who decide to become homeowners because of these loans are low-income households. Thus, mortgage contracts can influence asset decisions over the life cycle.

The next example considers the choice between a standard FRM and a CL when 80 percent of the home value is financed with a traditional fixed-payment mortgage and the other 20 percent is financed with another fixed-payment mortgage with a 2 percent interest rate premium. The aggregate homeownership rate in this economy is 68.65 percent. The introduction of a CL increases the homeownership rate by 4.3 percent. Table 15 shows how the homeownership rate decreases by age in this situation after the youngest cohort period. Just as with an ARM, the homeownership rate of the youngest cohort increases from 37 to nearly 43 percent. However, because the payments of the typical CL combo loan are higher than a corresponding ARM, income constraints prevent some young households from entering the market. Unlike the ARM, CL appears to have a positive effect across the entire age profile. Every age cohort has a homeownership rate at or above that in the baseline case.

The average home size in this economy is 1,909 square feet. This fact implies that the CL encourages the purchase of larger homes, which are affordable only for higher-income households. For this group, only 45.3 percent of the households have mortgage debt; 32.6 percent have a fixed-payment mortgage, and 12.7 percent have a CL. In addition, CLs are used in the bottom 40 percent of the income distribution. The income of an ARM household is lower than that of the average CL household.

CONCLUSION

This paper addresses several issues facing mortgage finance and potential home buyers. Recent innovations in the mortgage market have greatly expanded the types of loans available to home buyers. These products vary greatly in terms of payment size, composition of interest versus principal, and amortization schedule. Some products, such as interest-only loans, increase affordability by reducing payment size. However, these products typically slow accumulation of equity and thus become less attractive for wealth accumulation. Some mortgage types can generate negative amortization, which would seem highly unattractive to potential mortgage lenders. Other products, such as CLs, seek to increase affordability by reducing down payment requirements. These mortgages are characterized by larger mortgage payments. Given the typical government stance of seeking greater homeownership, both types of products appear successful in this regard.

In a standard macroeconomic model, we find that the typical ARM should generate large increases in the homeownership rate of young households. However, because of a delay in capital asset accumulation, lower homeownership may be found for older households. CLs also tend to drive up homeownership. For young households this increase in homeownership is not as pronounced as with ARMs, but with no apparent reduction in homeownership later in the life cycle. Thus, it should come as no surprise that the introduction of these mortgage products coincided with the observed increase in homeownership from 1995 through 2005. It should also not be surprising that the homeownership rate declines as these instruments are removed from the mortgage market.

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Real Interest Rate Persistence: Evidence and Implications

Christopher J. Neely and David E. Rapach

The real interest rate plays a central role in many important financial and macroeconomic models, including the consumption-based asset pricing model, neoclassical growth model, and models of the monetary transmission mechanism. The authors selectively survey the empirical literature that examines the time-series properties of real interest rates. A key stylized fact is that postwar real interest rates exhibit substantial persistence, shown by extended periods when the real interest rate is substantially above or below the sample mean. The finding of persistence in real interest rates is pervasive, appearing in a variety of guises in the literature. The authors discuss the implications of persistence for theoretical models, illustrate existing findings with updated data, and highlight areas for future research. (JEL C22, E21, E44, E52, E62, G12)

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The real interest rate—an interest rate adjusted for either realized or expected inflation—is the relative price of consuming now rather than later.[1] As such, it is a key variable in important theoretical models in finance and macroeconomics, such as the consumption-based asset pricing model (Lucas, 1978; Breeden, 1979; Hansen and Singleton, 1982, 1983), neoclassical growth model (Cass, 1965; Koopmans, 1965), models of central bank policy (Taylor, 1993), and numerous models of the monetary transmission mechanism.

The theoretical importance of the real interest rate has generated a sizable literature that examines its long-run properties. This paper selectively reviews this literature, highlights its central findings, and analyzes their implications for theory. We illustrate our study with new empirical results based on U.S. data. Two themes emerge from our review: (i) Real rates are very persistent, much more so than consumption growth; and (ii) researchers should seriously explore the causes of this persistence.

First, empirical studies find that real interest rates exhibit substantial *persistence*, shown by extended periods when postwar real interest rates are substantially above or below the sample mean. Researchers characterize this feature of the data with several types of models. One group of studies uses unit root and cointegration tests to analyze whether shocks permanently affect the real interest rate—that is, whether the real rate behaves like a random walk. Such studies often report evidence

[1] Heterogeneous agents face different real interest rates, depending on horizon, credit risk, and other factors. And inflation rates are not unique, of course. For ease of exposition, this paper ignores such differences as being irrelevant to the economic inference.

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of unit roots, or—at a minimum—substantial persistence. Other studies extend standard unit root and cointegration tests by considering whether real interest rates are fractionally integrated or exhibit significant nonlinear behavior, such as threshold dynamics or nonlinear cointegration. Fractional integration tests typically indicate that real interest rates revert to their mean very slowly. Similarly, studies that find evidence of nonlinear behavior in real interest rates identify regimes in which the real rate behaves like a unit root process. Another important group of studies reports evidence of structural breaks in the means of real interest rates. Allowing for such breaks reduces the persistence of deviations from the regime-specific means, so breaks reduce local persistence. The structural breaks themselves, however, still produce substantial global persistence in real interest rates.

The empirical literature thus finds that persistence is pervasive. Although researchers have used sundry approaches to model persistence, certain approaches are likely to be more useful than others. Comprehensive model selection exercises are thus an important area for future research, as they will illuminate the exact nature of real interest rate persistence.

The second theme of our survey is that the literature has not adequately addressed the *economic causes* of persistence in real interest rates. Understanding such processes is crucial for assessing the relevance of different theoretical models. We discuss potential sources of persistence and argue that monetary shocks contribute to persistent fluctuations in real interest rates. While identifying economic structure is always challenging, exploring the underlying causes of real interest rate persistence is an especially important area for future research.

The rest of the paper is organized as follows. The next section reviews the predictions of economic and financial models for the long-run behavior of the real interest rate. This informs our discussion of the theoretical implications of the empirical literature's results. After distinguishing between ex ante and ex post measures of the real interest rate, the third section reviews papers that apply unit root, cointegration, fractional integration, and nonlinearity tests to real interest rates. The fourth section discusses studies of regime switching and structural breaks in real interest rates. The fifth section considers sources of the persistence in the U.S. real interest rate and ultimately argues that it is a monetary phenomenon. The sixth section summarizes our findings.

THEORETICAL BACKGROUND

Consumption-Based Asset Pricing Model

The canonical consumption-based asset pricing model of Lucas (1978), Breeden (1979), and Hansen and Singleton (1982, 1983) posits a representative household that chooses a real consumption sequence, $\{c_t\}_{t=0}^{\infty}$, to maximize

$$\sum\nolimits_{t=0}^{\infty} \beta^t u(c_t),$$

subject to an intertemporal budget constraint, where β is a discount factor and $u(c_t)$ is an instantaneous utility function. The first-order condition leads to the familiar intertemporal Euler equation,

$$(1) \qquad E_t\left\{\beta\left[u'(c_{t+1})/u'(c_t)\right](1+r_t)\right\}=1,$$

where $1 + r_t$ is the gross one-period real interest rate (with payoff at period $t + 1$) and E_t is the conditional expectation operator. Researchers often assume that the utility function is of the constant relative risk aversion form, $u(c_t) = c_t^{1-\alpha}/(1-\alpha)$, where α is the coefficient of relative risk aversion. Combining this with the assumption of joint lognormality of consumption growth and the real interest rate implies the log-linear version of the first-order condition given by equation (1) (Hansen and Singleton, 1982, 1983):

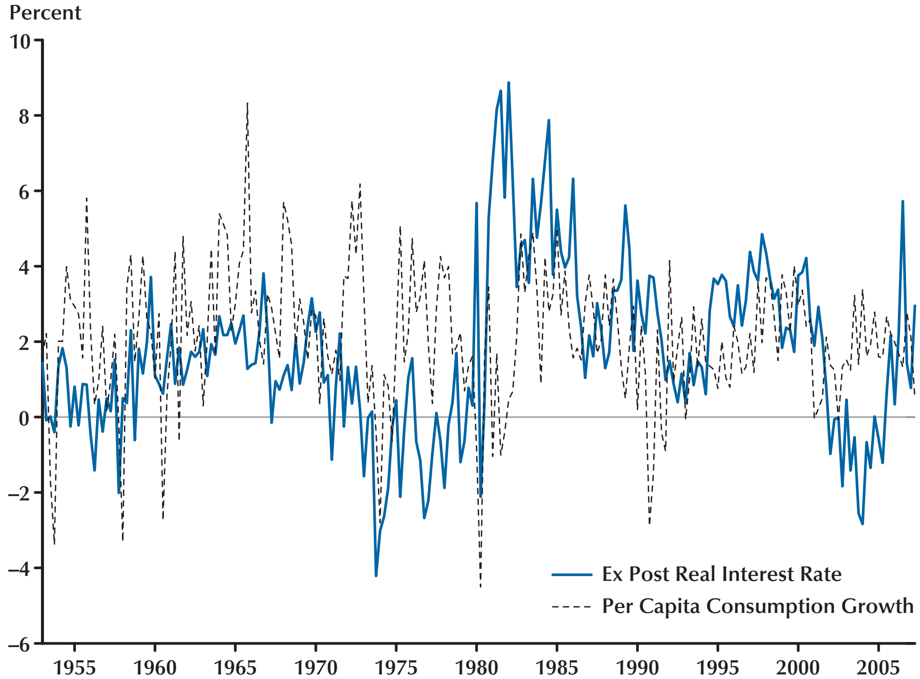
$$(2) \qquad \kappa-\alpha E_t\left[\Delta\log(c_{t+1})\right]+E_t\left[\log(1+r_t)\right]=0,$$

where $\Delta\log(c_{t+1}) = \log(c_{t+1}) - \log(c_t)$, $\kappa = \log(\beta) + 0.5\sigma^2$, and σ^2 is the constant conditional variance of $\log[\beta(c_{t+1}/c_t)^{-\alpha}(1 + r_t)]$.

Equation (2) links the conditional expectations of the growth rate of real per capita consumption $[\Delta\log(c_{t+1})]$ with the (net) real interest rate $[\log(1 + r_t) \cong r_t]$. Rose (1988) argues that if equation (2) is to hold, then these two series must have

Figure 1

U.S. Ex Post Real Interest Rate and Real Per Capita Consumption Growth, 1953:Q1–2007:Q2



NOTE: The figure plots the U.S. ex post 3-month real interest rate and annualized per capita consumption growth. Consumption is measured as the sum of nondurable goods and services consumption.

similar integration properties. Whereas $\Delta\log(c_{t+1})$ is almost surely a stationary process [$\Delta\log(c_{t+1}) \sim I(0)$], Rose (1988) presents evidence that the real interest rate contains a unit root [$r_t \sim I(1)$] in many industrialized countries. A unit root in the real interest rate combined with stationary consumption growth means that there will be permanent changes in the level of the real rate not matched by such changes in consumption growth, so equation (2) apparently cannot hold.

Figure 1 illustrates the problem identified by Rose (1988) using U.S. data for the ex post 3-month real interest rate and annualized growth rate of per capita consumption (nondurable goods plus services) for 1953:Q1–2007:Q2. The two series appear to track each other reasonably well for long periods, such as the 1950s, 1960s, and 1984-2001, but they also diverge for significant periods, such as the 1970s, early 1980s, and 2001-05.

The simplest versions of the consumption-based asset pricing model are based on an endowment economy with a representative household and constant preferences. The next subsection discusses the fact that more elaborate theoretical models allow for some changes in the economy—for example, changes in fiscal or monetary policy—to alter the steady-state real interest rate while leaving steady-state consumption growth unchanged. That is, they permit a mismatch in the integration properties of the real interest rate and consumption growth.

Equilibrium Growth Models and the Steady-State Real Interest Rate

General equilibrium growth models with a production technology imply Euler equations similar to equations (1) and (2) that suggest sources of a unit root in real interest rates. Specifically, the

Cass (1965) and Koopmans (1965) neoclassical growth model with a representative profit-maximizing firm and utility-maximizing household predicts that the steady-state real interest rate is a function of time preference, risk aversion, and the steady-state growth rate of technological change (Blanchard and Fischer, 1989, Chap. 2; Barro and Sala-i-Martin, 2003, Chap. 3; Romer, 2006, Chap. 2). In this model the assumption of constant relative risk aversion utility implies the following familiar steady-state condition:

$$r^* = \zeta + \alpha z, \tag{3}$$

where r^* is the steady-state real interest rate, $\zeta = -\log(\beta)$ is the rate of time preference, and z is the (expected) steady-state growth rate of labor-augmenting technological change. Equation (3) implies that a permanent change in the exogenous rate of time preference, risk aversion, or long-run growth rate of technology will affect the steady-state real interest rate.[2] If there is no uncertainty, the neoclassical growth model implies the following steady-state version of the Euler equation given by (2):

$$-\zeta - \alpha\left[\Delta \log(c)\right]^* + r^* = 0, \tag{4}$$

where $[\Delta\log(c)]^*$ represents the steady-state growth rate of c_t. Substituting the right-hand side of equation (3) into equation (4) for r^*, one finds that steady-state technology growth determines steady-state consumption growth: $[\Delta\log(c)]^* = z$.

If the rate of time preference (ζ), risk aversion (α), and/or steady-state rate of technology growth (z) change, then (3) requires corresponding changes in the steady-state real interest rate. Depending on the size and frequency of such changes, real interest rates might be very persistent, exhibiting unit root behavior and/or structural breaks. Of these three factors, a change in the steady-state growth rate of technology—such as those that might be associated with the "productivity slowdown" of the early 1970s and/or the "New Economy" resurgence of the mid-1990s—is the only one that will alter both the real interest rate and consumption growth, producing nonstationary behavior in both variables. Thus, it cannot explain the mismatch in the integration properties of the real interest rate and consumption growth identified by Rose (1988).

On the other hand, shocks to the preference parameters, ζ and α, will change only the steady-state real interest rate and not steady-state consumption growth. Therefore, changes in preferences potentially disconnect the integration properties of real interest rates and consumption growth. Researchers generally view preferences as stable, however, making it unpalatable to ascribe the persistence mismatch to such changes.[3]

In more elaborate models, still other factors can change the steady-state real interest rate. For example, permanent changes in government purchases and their financing can also affect the steady-state real rate in overlapping generations models with heterogeneous households (Samuelson, 1958; Diamond, 1965; Blanchard, 1985; Blanchard and Fischer, 1989, Chap. 3; Romer, 2006, Chap. 2). Such shocks affect the steady-state real interest rate without affecting steady-state consumption growth, so they potentially explain the mismatch in the integration properties of the real interest rate and consumption growth examined by Rose (1988).

Finally, some monetary growth models allow for changes in steady-state money growth to affect the steady-state real interest rate. The seminal models of Mundell (1963) and Tobin (1965) predict that an increase in steady-state money growth lowers the steady-state real interest rate, and more recent micro-founded monetary models have similar implications (Weiss, 1980; Espinosa-Vega and Russell, 1998a,b; Bullard and Russell, 2004; Reis, 2007; Lioui and Poncet, 2008). Again, this class of models permits changes in the steady-state real interest rate without corresponding changes in consumption growth, potentially explaining a mismatch in the integration properties of the real interest rate and consumption growth.

[2] Changes in distortionary tax rates could also affect r^* (Blanchard and Fischer, 1989, pp. 56-59).

[3] Some researchers appear more willing to allow for changes in preferences over an extended period. For example, Clark (2007) argues that a steady decrease in the rate of time preference is responsible for the downward trend in real interest rates in Europe from the early medieval period to the eve of the Industrial Revolution.

Transitional Dynamics

The previous section discusses factors that can affect the *steady-state* real interest rate. Other shocks can have persistent—but ultimately transitory—effects on the real rate. For example, in the neoclassical growth model, a temporary increase in technology growth or government purchases leads to a persistently (but not permanently) higher real interest rate (Romer, 2006, Chap. 2). In addition, monetary shocks can persistently affect the real interest rate via a variety of frictions, such as "sticky" prices and information, adjustment costs, and learning by agents about policy regimes. Transient technology and fiscal shocks, as well as monetary shocks, can also explain differences in the persistence of real interest rates and consumption growth. For example, using a calibrated neoclassical equilibrium growth model, Baxter and King (1993) show that a temporary (four-year) increase in government purchases persistently raises the real interest rate, although it eventually returns to its initial level. In contrast, the fiscal shock produces a much less persistent reaction in consumption growth. As we will discuss later, evidence of highly persistent but mean-reverting behavior in real interest rates supports the empirical relevance of these shocks.

TESTING THE INTEGRATION PROPERTIES OF REAL INTEREST RATES

Ex Ante versus Ex Post Real Interest Rates

The ex ante real interest rate (EARR) is the nominal interest rate minus the expected inflation rate, while the ex post real rate (EPRR) is the nominal rate minus actual inflation. Agents make economic decisions on the basis of their inflation expectations over the decision horizon. For example, the Euler equations (1) and (2) relate the expected marginal utility of consumption to the expected real return. Therefore, the EARR is the relevant measure for evaluating economic decisions, and we really wish to evaluate the EARR's time-series properties, rather than those of the EPRR.

Unfortunately, the EARR is not directly observable because expected inflation is not directly observable. An obvious solution is to use some survey measure of inflation expectations, such as the Livingston Survey of professional forecasters, which has been conducted biannually since the 1940s (Carlson, 1977). Economists are often reluctant, however, to accept survey forecasts as expectations. For example, Mishkin (1981, p. 153) expresses "serious doubts as to the quality of these [survey] data." Obtaining survey data at the desired frequency for the desired sample might create other obstacles to the use of survey data. Some studies have used survey data, however, including Crowder and Hoffman (1996) and Sun and Phillips (2004).

There are at least two alternative approaches to the problem of unobserved expectations. The first is to use econometric forecasting methods to construct inflation forecasts; see, for example, Mishkin (1981, 1984) and Huizinga and Mishkin (1986). Unfortunately, econometric forecasting models do not necessarily include all of the relevant information agents use to form expectations of inflation, and such models can fail to change with the structure of the economy. For example, Stock and Watson (1999, 2003) show that both real activity and asset prices forecast inflation but that the predictive relations change over time.[4]

A second alternative approach is to use the actual inflation rate as a proxy for inflation expectations. By definition, the actual inflation rate at time t (π_t) is the sum of the expected inflation rate and a forecast error term (ε_t):

$$\pi_t = E_{t-1}\pi_t + \varepsilon_t. \tag{5}$$

The literature on real interest rates has long argued that, if expectations are formed rationally, $E_{t-1}\pi_t$ should be an optimal forecast of inflation (Nelson and Schwert, 1977), and ε_t should there-

[4] Atkeson and Ohanian (2001) and Stock and Watson (2007) discuss the econometric challenges in forecasting inflation. One might also consider using Treasury inflation-protected securities (TIPS) yields—and/or their foreign counterparts—to measure real interest rates. But these series have a relatively short span of available data, in that the U.S. securities were first issued in 1997, are only available at long maturities (5, 10, and 20 years), and do not correctly measure real rates when there is a significant chance of deflation.

fore be a white noise process. The EARR can be expressed (approximately) as

$$r_t^{ea} = i_t - E_t \pi_{t+1}, \tag{6}$$

where i_t is the nominal interest rate. Solving equation (5) for $E_t(\pi_{t+1})$ and substituting it into equation (6), we have

$$\begin{aligned} r_t^{ea} &= i_t - \left(\pi_{t+1} - \varepsilon_{t+1}\right) \\ &= i_t - \pi_{t+1} + \varepsilon_{t+1} = r_t^{ep} + \varepsilon_{t+1}, \end{aligned} \tag{7}$$

where $r_t^{ep} = i_t - \pi_{t+1}$ is the EPRR. Equation (7) implies that, under rational expectations, the EPRR and EARR differ only by a white noise component, so the EPRR and EARR will share the same long-run (integration) properties. Actually, this latter result does not require expectations to be formed rationally but holds if the expectation errors (ε_{t+1}) are stationary.[5] Beginning with Rose (1988), much of the empirical literature tests the integration properties of the EARR with the EPRR, after assuming that inflation-expectation errors are stationary.

Researchers typically evaluate the integration properties of the EPRR with a decision rule. They first analyze the individual components of the EPRR, i_t and π_{t+1}. If unit root tests indicate that i_t and π_{t+1} are both $I(0)$, then this implies a stationary EPRR, as any linear combination of two $I(0)$ processes is also an $I(0)$ process.[6] If i_t and π_{t+1} have different orders of integration—for example, if $i_t \sim I(1)$ and $\pi_{t+1} \sim I(0)$—then the EPRR must have a unit root, as any linear combination of an $I(1)$ process and an $I(0)$ process is an $I(1)$ process. Finally, if unit root tests show that i_t and π_{t+1} are both $I(1)$, researchers test for a stationary EPRR by testing for cointegration between i_t and π_{t+1}—that is, testing whether the linear combination $i_t - [\theta_0 + \theta_1 \pi_{t+1}]$ is a stationary process—using one of two approaches.[7] First, many researchers impose a cointegrating vector of $(1,-\theta_1)' = (1,-1)'$ and apply unit root tests to $r_t^{ep} = i_t - \pi_{t+1}$. This approach typically has more power to reject the null of no cointegration when the true cointegrating vector is $(1,-1)'$. The second approach is to freely estimate the cointegrating vector between i_t and π_{t+1}, as this allows for tax effects (Darby, 1975).

If $i_t, \pi_{t+1} \sim I(1)$, then a stationary EPRR requires i_t and π_{t+1} to be cointegrated with cointegrating coefficient, $\theta_1 = 1$, or, allowing for tax effects, $\theta_1 = 1/(1-\tau)$, where τ is the marginal investor's marginal tax rate on nominal interest income. When allowing for tax effects, researchers view estimates of θ_1 in the range of 1.3 to 1.4 as plausible, as they correspond to a marginal tax rate around 0.2 to 0.3 (Summers, 1983).[8] It is worth emphasizing that cointegration between i_t and π_{t+1} by itself does not imply a stationary real interest rate: θ_1 must also equal 1 [or $1/(1-\tau)$], as other values of θ_1 imply that the equilibrium real interest rate varies with inflation.

Although much of the empirical literature analyzes the EPRR in this manner, it is important to keep in mind that the EPRR's time-series properties can differ from those of the EARR—the ultimate object of analysis—in two ways. First, the EPRR's behavior at short horizons might differ from that of the EARR. For example, using survey data and various econometric methods to forecast inflation, Dotsey, Lantz, and Scholl (2003) study the behavior of the EARR and EPRR at business-cycle frequencies and find that their behavior over the business cycle can differ significantly. Second, some estimation techniques can generate different persistence properties between the EARR and EPRR; see, for example, Evans and Lewis (1995) and Sun and Phillips (2004).

Early Studies

A collection of early studies on the efficient market hypothesis and the ability of nominal

[5] Peláez (1995) provides evidence that inflation-expectation errors are stationary. Also note that Andolfatto, Hendry, and Moran (2008) argue that inflation-expectation errors can appear serially correlated in finite samples, even when expectations are formed rationally, due to short-run learning dynamics about infrequent changes in the monetary policy regime.

[6] The appendix, "Unit Roots and Cointegration Tests," provides more information on the mechanics of popular unit root and cointegration tests.

[7] The presence of θ_0 allows for a constant term in the cointegrating relationship corresponding to the steady-state real interest rate.

[8] Data from tax-free municipal bonds would presumably provide a unitary coefficient. Crowder and Wohar (1999) study the Fisher effect with tax-free municipal bonds.

interest rates to forecast the inflation rate foreshadows the studies that use unit root and cointegration tests. Fama (1975) presents evidence that the monthly U.S. EARR can be viewed as constant over 1953-71. Nelson and Schwert (1977), however, argue that statistical tests of Fama (1975) have low power and that his data are actually not very informative about the EARR's autocorrelation properties. Hess and Bicksler (1975), Fama (1976), Carlson (1977), and Garbade and Wachtel (1978) also challenge Fama's (1975) finding on statistical grounds. In addition, subsequent studies show that Fama's (1975) result hinges critically on the particular sample period (Mishkin, 1981, 1984; Huizinga and Mishkin, 1986; Antoncic, 1986).

Unit Root and Cointegration Tests

The development of unit root and cointegration analysis, beginning with Dickey and Fuller (1979), spurred the studies that formally test the persistence of real interest rates. In his seminal study, Rose (1988) tests for unit roots in short-term nominal interest rates and inflation rates using monthly data for 1947-86 for 18 countries in the Organisation for Economic Co-operation and Development (OECD). Rose (1988) finds that augmented Dickey-Fuller (ADF) tests fail to reject the null hypothesis of a unit root in short-term nominal interest rates, but they can consistently reject a unit root in inflation rates based on various price indices—consumer price index (CPI), gross national product (GNP) deflator, implicit price deflator, and wholesale price index (WPI).[9] As discussed above, the finding that $i_t \sim I(1)$ while $\pi_t \sim I(0)$ indicates that the EPRR, $i_t - \pi_{t+1}$, is an $I(1)$ process. Under the assumption that inflation-expectation errors are stationary, this also implies that the EARR is an $I(1)$ process. Rose (1988) easily rejects the unit root null hypothesis for U.S. consumption growth, which leads him to argue that an $I(1)$ real interest rate and $I(0)$ consumption growth rate violates the intertemporal Euler equation implied by the consumption-based asset pricing model. Beginning with Rose (1988), Table 1 summarizes the methods and conclusions of surveyed papers on the long-run properties of real interest rates.

A number of subsequent papers also test for a unit root in real interest rates. Before estimating structural vector autoregressive (SVAR) models, King et al. (1991) and Galí (1992) apply ADF unit root tests to the U.S. nominal 3-month Treasury bill rate, inflation rate, and EPRR. Using quarterly data for 1954-88 and the GNP deflator inflation rate, King et al. (1991) fail to reject the null hypothesis of a unit root in the nominal interest rate, matching the finding of Rose (1988). Unlike Rose (1988), however, King et al. cannot reject the unit root null hypothesis for the inflation rate, which creates the possibility that the nominal interest rate and inflation rate are cointegrated. Imposing a cointegrating vector of (1,–1)′, they fail to reject the unit root null hypothesis for the EPRR. Using quarterly data for 1955-87, the CPI inflation rate, and simulated critical values that account for potential size distortions due to moving-average components, Galí (1992) obtains unit root test results similar to those of King et al. Despite the failure to reject the null hypothesis that $i_t - \pi_{t+1} \sim I(1)$, Galí nevertheless assumes that $i_t - \pi_{t+1} \sim I(0)$ when he estimates his SVAR model, contending that "the assumption of a unit root in the real [interest] rate seems rather implausible on a priori grounds, given its inconsistency with standard equilibrium growth models" (Galí, 1992, p. 717). This is in interesting contrast to King et al., who maintain the assumption that $i_t - \pi_{t+1} \sim I(1)$ in their SVAR model. Shapiro and Watson (1988) report similar unit root findings and, like Galí, still assume the EPRR is stationary in an SVAR model.

Analyzing a 1953-90 full sample, as well as a variety of subsamples for the nominal Treasury bill rate and CPI inflation rate, Mishkin (1992) argues that monthly U.S. data are largely consistent with a stationary EPRR. With simulated critical values, as in Galí (1992), Mishkin (1992) finds that the nominal interest rate and inflation rate are both $I(1)$ over four sample periods: 1953:01–1990:12, 1953:01–1979:10, 1979:11–1982:10, and 1982:11–1990:12. He then tests whether the nominal interest rate and inflation rate are cointegrated using both the single-equation augmented Engle and Granger (1987, AEG) test and by prespecify-

[9] The appendix discusses unit root and cointegration tests.

Table 1

Selective Summary of the Empirical Literature on the Long-Run Properties of Real Interest Rates

Study	Sample	Countries	Nominal interest rate and price data
Rose (1988)	A: 1892-70, 1901-50 Q: 1947-86 M: 1948-86	18 OECD countries	Long-term corporate bond yield, short-term commercial paper rate, GNP deflator, CPI, implicit price deflator, WPI
King et al. (1991)	Q: 1949-88	U.S.	3-month U.S. Treasury bill rate, implicit GNP deflator
Galí (1992)	Q: 1955-87	U.S.	3-month U.S. Treasury bill rate, CPI
Mishkin (1992)	M: 1953-90	U.S.	1- and 3-month Treasury bill rates, CPI
Wallace and Warner (1993)	Q: 1948-90	U.S.	3-month Treasury bill rate, 10-year government bond yield, CPI
Engsted (1995)	Q: 1962-93	13 OECD countries	Long-term bond yield, CPI
Mishkin and Simon (1995)	Q: 1962-93	Australia	13-week government bond yield, CPI
Crowder and Hoffman (1996)	Q: 1952-91	U.S.	3-month Treasury bill rate, implicit consumption deflator, Livingston inflation expectations survey, tax data from various sources
Koustas and Serletis (1999)	Q: Data begin from 1957-72; all data end in 1995	11 OECD countries	Various short-term nominal interest rates, CPI
Bierens (2000)	M: 1954-94	U.S.	Federal funds rate, CPI
Rapach (2003)	A: Data begin in 1949-65; end in 1994-96	14 industrialized countries	Long-term government bond yield, implicit GDP deflator
Rapach and Weber (2004)	Q: 1957-2000	16 OECD countries	Long-term government bond yield, CPI
Rapach and Wohar (2004)	Q: 1960-1998	13 OECD countries	Long-term government bond yield, CPI marginal tax rate data (Padovano and Galli, 2001)

NOTE: A, Q, and M indicate annual, quarterly, and monthly data frequencies; GNP denotes gross national product.

Results on the long-run properties of nominal interest rates, inflation rates, and real interest rates
ADF tests fail to reject a unit root for nominal interest rates but do reject for inflation rates, indicating a unit root in EPRRs. ADF tests do reject a unit root for consumption growth.
ADF tests fail to reject a unit root for the nominal interest rate, inflation rate, and EPRR.
ADF tests with simulated critical values that adjust for moving-average components fail to reject a unit root in the nominal interest rate, inflation rate, and EPRR.
ADF tests with simulated critical values that adjust for moving-average components fail to reject a unit root in the nominal interest rate and inflation rate. AEG tests typically reject the null of no cointegration, indicating a stationary EPRR.
ADF tests fail to reject a unit root in the long-term nominal interest rate and inflation rate. Johansen (1991) procedure provides evidence that the variables are cointegrated and that the EPRR is stationary.
ADF tests fail to reject a unit root in nominal interest rates and inflation rates, while cointegration tests present ambiguous results on the stationarity of the EPRR across countries.
ADF tests fail to reject a unit root in the nominal interest rate and inflation rate. AEG tests typically fail to reject the null hypothesis of no cointegration, indicating a nonstationary EPRR.
ADF test fails to reject a unit root in the nominal interest rate and inflation rate after accounting for moving-average components. Johansen (1991) procedure rejects the null of no cointegration and supports a stationary EPRR.
ADF tests usually fail to reject a unit root in nominal interest rates and inflation rates, while KPSS tests typically reject the null of stationarity, indicating nonstationary nominal interest rates and inflation rates. AEG tests typically fail to reject the null of no cointegration, indicating a nonstationary EPRR.
New test provides evidence of nonlinear cotrending between the nominal interest rate and inflation rate, indicating a stationary EPRR. New test, however, cannot distinguish between nonlinear cotrending and linear cointegration.
ADF tests fail to reject a unit root in all nominal interest rates and in 13 of 17 inflation rates. This indicates a nonstationary EPRR for the four countries with a stationary inflation rate. AEG tests typically fail to reject a unit root in the EPRR for the 13 countries with a nonstationary inflation rate, indicating a nonstationary EPRR for these countries.
Ng and Perron (2001) unit root tests typically fail to reject a unit root in nominal interest rates and inflation rates. Ng and Perron (2001) and Perron and Rodriguez (2001) tests usually fail to reject the null of no cointegration, indicating a nonstationary EPRR in most countries.
Lower (upper) 95 percent confidence band for the EPRR's ρ is close to 0.90 (above unity) for nearly every country.

Table 1, cont'd

Selective Summary of the Empirical Literature on the Long-Run Properties of Real Interest Rates

Study	Sample	Countries	Nominal interest rate and price data
Karanasos, Sekioua, and Zeng (2006)	A: 1876-2000	U.S.	Long-term government bond yield, CPI
Lai (1997)	Q: 1974-2001	8 industrialized and 8 developing countries	1- to 12-month Treasury bill rates, CPI, Data Resources, Inc. inflation forecasts
Tsay (2000)	M: 1953-90	U.S.	1- and 3-month Treasury bill rates, CPI
Sun and Phillips (2004)	Q: 1934-94	U.S.	3-month Treasury bill rate, inflation forecasts from the *Survey of Professional Forecasters*, CPI
Pipatchaipoom and Smallwood (2008)	M: 1971-2003	U.S.	Eurodollar rate, CPI
Maki (2003)	M: 1972-2000	Japan	10-year bond rate, call rate, CPI
Million (2004)	M: 1951-99	U.S.	3-month Treasury bill rate, CPI
Christopoulos and León-Ledesma (2007)	Q: 1960-2004	U.S.	3-month Treasury bill rate, CPI
Koustas and Lamarche (2008)	A: 1960-2004	G-7 countries	3-month government bill rate, CPI
Garcia and Perron (1996)	Q: 1961-86	U.S.	3-month Treasury bill rate, CPI
Clemente, Montañés, and Reyes (1998)	Q: 1980-95	U.S., U.K.	Long-term government bond yield, CPI
Caporale and Grier (2000)	Q: 1961-86	U.S.	3-month Treasury bill rate, CPI
Bai and Perron (2003)	Q: 1961-86	U.S.	3-month Treasury bill rate, CPI

NOTE: A, Q, and M indicate annual, quarterly, and monthly data frequencies; GNP denotes gross national product.

Results on the long-run properties of nominal interest rates, inflation rates, and real interest rates
95 percent confidence interval for the EPRR's ρ is (0.97, 0.99). There is evidence of long-memory, mean-reverting behavior in the EPRR.
ADF and KPSS tests indicate a unit root in the nominal interest rate, inflation rate, and expected inflation rate. There is evidence of long-memory, mean-reverting behavior in the EARR and EPRR.
There is evidence of long-memory, mean-reverting behavior in the EPRR.
Bivariate exact Whittle estimator indicates long-memory behavior in the EARR. There is no evidence of a fractional cointegrating relationship between the nominal interest rate and expected inflation rate.
Exact Whittle estimator provides evidence of long-memory, mean-reverting behavior in the EARR.
Breitung (2002) nonparametric test that allows for nonlinear short-run dynamics provides evidence of cointegration between the nominal interest rate and inflation rate; cointegrating vector is not estimated, however, so it is not known if the cointegrating relationship is consistent with a stationary EPRR.
Luukkonen, Saikkonen, and Teräsvirta (1988) test rejects linear short-run dynamics for the adjustment to the long-run equilibrium EPRR. A smooth transition autoregressive model exhibits asymmetric mean reversion in the EPRR, depending on the level of the EPRR.
Choi and Saikkonen (2005) test provides evidence of nonlinear cointegration between the nominal interest rate and inflation rate. Exponential smooth transition regression (ESTR) model fits best over the full sample and the first subsample (1960-78), while a logistic smooth transition regression (LSTR) model fits best over the second subsample (1979-2004). Estimated ESTR model for 1960-78 is not consistent with a stationary EPRR for any inflation rate, and estimated LSTR model for 1979–2004 is consistent with a stationary EPRR only when the inflation rate is above approximately 3 percent.
ADF and KPSS tests provide evidence of a unit root in the nominal interest rate and inflation rate. Bec, Ben Salem, and Carassco (2004) nonlinear unit root and Hansen (1996, 1997) linearity tests indicate that the EPRR can be suitably modeled as a three-regime self-exciting autoregressive (SETAR) process in Canada, France, and Italy.
An estimated autoregressive model with a three-state Markov-switching process for the mean indicates that the EPRR was in a "moderate"-mean regime for 1961-73, a "low"-mean regime for 1973-80, and a "high"-mean regime for 1980-86. EPRR is stationary with little persistence within these regimes.
ADF tests that allow for two structural breaks in the mean reject a unit root in the EPRR, indicating that the EPRR is stationary within regimes defined by structural breaks.
Bai and Perron (1998) methodology provides evidence of multiple structural breaks in the mean EPRR.
Bai and Perron (1998) methodology provides evidence of multiple structural breaks in the mean EPRR.

Table 1, cont'd

Selective Summary of the Empirical Literature on the Long-Run Properties of Real Interest Rates

Study	Sample	Countries	Nominal interest rate and price data
Lai (2004)	M: 1978-2002	U.S.	1-year Treasury bill rate, inflation expectations from the University of Michigan Survey of Consumers, CPI, federal marginal income tax rates for four-person families
Rapach and Wohar (2005)	Q: 1960-98	13 OECD countries	Long-term government bond yield, CPI, marginal tax rate data (Padovano and Galli, 2001)
Lai (2008)	Q: 1974-2001	8 industrialized and 8 developing countries	1- to 12-month Treasury bill rate, deposit rate, CPI

NOTE: A, Q, and M indicate annual, quarterly, and monthly data frequencies; GNP denotes gross national product.

ing a cointegrating vector and testing for a unit root in $i_t - \pi_{t+1}$. Mishkin (1992) rejects the null hypothesis of no cointegration for the 1953:01–1990:12 and 1953:01–1979:10 periods, but finds less frequent and weaker rejections for the 1979:11–1982:10 and 1982:11–1990:12 periods.[10] Mishkin and Simon (1995) apply similar tests to quarterly short-term nominal interest rate and inflation rate data for Australia. Using a 1962:Q3–1993:Q4 full sample, as well as 1962:Q3–1979:Q3 and 1979:Q4– 1993:Q4 subsamples, they find evidence that both the nominal interest rate and the inflation rate are $I(1)$, agreeing with the results for U.S. data in Mishkin (1992). There is weaker evidence that the Australian nominal interest rate and inflation rate are cointegrated than there is for U.S. data. Nevertheless, Mishkin and Simon (1995) argue that theoretical considerations warrant viewing the long-run real interest rate as stationary in Australia, as "any reasonable model of the macro economy would surely suggest that real interest rates have mean-reverting tendencies which make them stationary" (Mishkin and Simon, 1995, p. 223).

Koustas and Serletis (1999) test for unit roots and cointegration in short-term nominal interest rates and CPI inflation rates using quarterly data for 1957-95 for 11 industrialized countries. They use ADF unit root tests as well as the KPSS unit root test of Kwiatkowski et al. (1992), which takes stationarity as the null hypothesis and nonstationarity as the alternative. ADF and KPSS unit root tests indicate that $i_t \sim I(1)$ and $\pi_{t+1} \sim I(1)$ in most countries, so a stationary EPRR requires cointegration between the nominal interest rate and inflation rate. Koustas and Serletis (1999), however, usually fail to find strong evidence of cointegration using the AEG test. Overall, their study finds that the EPRR is nonstationary in many industrialized countries. Rapach (2003) obtains similar results using postwar data for an even larger number of OECD countries.

In a subtle variation on conventional cointegration analysis, Bierens (2000) allows an individual time series to have a deterministic component that is a highly complex function of time—essentially a smooth spline—and a stationary stochastic component, and he develops nonparametric procedures to test whether two series share a common

[10] Although they use essentially the same econometric procedures and similar samples, Galí (1992) is unable to reject the unit root null hypothesis for the EPRR, while Mishkin (1992) does reject this null hypothesis. This illustrates the sensitivity of EPRR unit root and cointegration tests to the specific sample. In addition, the use of short samples, such as the 1979:11–1982:10 sample period considered by Mishkin (1992), is unlikely to be informative about the integration properties of the EPRR. To infer long-run behavior, one needs reasonably long samples.

Results on the long-run properties of nominal interest rates, inflation rates, and real interest rates
ADF tests allowing for a structural break in the mean reject a unit root in the tax-adjusted or unadjusted EARR, indicating that the EARR is stationary within regimes defined by the structural break.
The Bai and Perron (1998) methodology provides evidence of structural breaks (usually multiple) in the mean EPRR and mean inflation rate for all 13 countries.
ADF tests allowing for a structural break in the mean reject a unit root in the EPRR for most countries, indicating that the EPRR is stationary within regimes defined by the structural break.

deterministic component ("nonlinear cotrending"). Using monthly U.S. data for 1954-94, Bierens (2000) presents evidence that the federal funds rate and CPI inflation rate cotrend with a vector of (1,–1)′, which can be interpreted as evidence for a stationary real interest rate. Bierens shows, however, that his tests cannot differentiate between nonlinear cotrending and linear cointegration in the presence of stochastic trends in the nominal interest rate and inflation rate. In essence, the highly complex deterministic components for the individual series closely mimic unit root behavior.

A number of studies use the Johansen (1991) system–based cointegration procedure to test for a stationary EPRR. Wallace and Warner (1993) apply the Johansen (1991) procedure to quarterly U.S. nominal 3-month Treasury bill rate and CPI inflation data for a 1948-90 full sample and a number of subsamples. Their results generally support the existence of a cointegrating relationship, and their estimates of θ_1 are typically not significantly different from unity, in line with a stationary EPRR. Wallace and Warner (1993) also argue that the expectations hypothesis implies that short-term and long-term nominal interest rates should be cointegrated, and they find evidence that U.S. short and long rates are cointegrated with a cointegrating vector of (1,–1)′. In line with the results for the nominal 3-month Treasury bill rate, Wallace and Warner find that the nominal 10-year Treasury bond rate and inflation rate are cointegrated.

With quarterly U.S. data for 1951-91, Crowder and Hoffman (1996) also use the Johansen (1991) procedure to test for cointegration between the 3-month Treasury bill rate and implicit consumption deflator inflation rate. As in Wallace and Warner (1993), they reject the null of no cointegration between the nominal interest rate and inflation rate. Their estimates of θ_1 range from 1.22 to 1.34, which are consistent with a stationary tax-adjusted EPRR. Crowder and Hoffman (1996) also use estimates of average marginal tax rates to directly test for cointegration between $i_t(1-\tau)$ and π_{t+1}. The Johansen (1991) procedure supports cointegration and estimates a cointegrating vector not significantly different from (1,–1)′, in line with a stationary tax-adjusted EPRR.

Engsted (1995) uses the Johansen (1991) procedure to test for cointegration between the nominal long-term government bond yield and CPI inflation rate in 13 OECD countries using quarterly data for 1962-93. In broad agreement with the results of Wallace and Warner (1993) and Crowder and Hoffman (1996), Engsted (1995) rejects the

Table 2

Unit Root Test Statistics, U.S. data, 1953:Q1–2007:Q2

Variable	ADF	MZ_α
3-Month Treasury bill rate	–2.49 [7]	–4.39 [8]
PCE deflator inflation rate	–2.72* [4]	–5.20 [5]
Ex post real interest rate	–3.06** [6]	–18.83*** [2]
Per capita consumption growth	–4.99*** [4]	–42.07*** [2]

NOTE: The ADF and MZ_α statistics correspond to a one-sided (lower-tail) test of the null hypothesis that the variable has a unit root against the alternative hypothesis that the variable is stationary. The 10 percent, 5 percent, and 1 percent critical values for the ADF statistic are –2.58, –2.89, and –3.51; the 10 percent, 5 percent, and 1 percent critical values for the MZ_α statistic are –5.70, –8.10, and –13.80. The lag order for the regression model used to compute the test statistic is reported in brackets. *, **, and *** indicate significance at the 10 percent, 5 percent, and 1 percent levels. PCE denotes personal consumption expenditures.

null hypothesis of no cointegration for almost all countries. The estimates of θ_1 vary quite markedly across countries, however, and the values are often inconsistent with a stationary EPRR.

Overall, unit root and cointegration tests present mixed results with respect to the integration properties of the EPRR. Generally speaking, single-equation methods provide weaker evidence of a stationary EPRR, while the Johansen (1991) system–based approach supports a stationary EPRR, at least for the United States. Unfortunately, econometric issues, such as the low power of unit root tests and size distortions in the presence of moving-average components, complicate inference about persistence.

To address these econometric issues, Rapach and Weber (2004) use unit root and cointegration tests with improved size and power. Specifically, they use the Ng and Perron (2001) unit root and Perron and Rodriguez (2001) cointegration tests. These tests incorporate aspects of the modified ADF tests in Elliott, Rothenberg, and Stock (1996) and Perron and Ng (1996), as well as an adjusted modified information criterion to select the autoregressive (AR) lag order, to develop tests that avoid size distortions while retaining power. Rapach and Weber (2004) use quarterly nominal long-term government bond yield and CPI inflation rate data for 1957-2000 for 16 industrialized countries. The Ng and Perron (2001) unit root and Perron and Rodriguez (2001) cointegration tests provide mixed results, but Rapach and Weber interpret their results as indicating that the EPRR is nonstationary in most industrialized countries over the postwar era.

Updated Unit Root and Cointegration Test Results for U.S. Data

Tables 2 and 3 illustrate the type of evidence provided by unit root and cointegration tests for the U.S. 3-month Treasury bill rate, CPI inflation rate, and per capita consumption growth rate for 1953:Q1–2007:Q2 (the same data as in Figure 1).

Table 2 reports the ADF statistic, as well as the MZ_α statistic from Ng and Perron (2001), which is designed to have better size and power properties than the former. Consistent with the literature, neither test rejects the unit root null hypothesis for the nominal interest rate. The results are mixed for the inflation rate: The ADF statistic rejects the unit root null at the 10 percent level, but the MZ_α statistic does not reject at conventional significance levels. The ADF test result that $i_t \sim I(1)$ while $\pi_t \sim I(0)$ means that the EPRR is nonstationary, as in Rose (1988).[11] The MZ_α statistic's failure to reject the unit root null for either inflation or nomi-

[11] A significant moving-average component in the inflation rate could create size distortions in the ADF statistic that lead us to falsely reject the unit root null hypothesis for that series. The fact that we do not reject the unit root null using the MZ_α statistic—which is designed to avoid this size distortion—supports this interpretation. Rapach and Weber (2004), however, do reject the unit root null for the U.S. inflation rate using the MZ_α statistic and data through 2000. Inflation rate unit root tests are thus particularly sensitive to the sample period.

Table 3

Cointegration Test Statistics and Cointegrating Coefficient Estimates, U.S. 3-Month Treasury Bill Rate and Inflation Rate (1953:Q1–2007:Q2)

Cointegration tests		
AEG	MZ_α	Trace
–3.07* [6]	–17.11** [2]	19.96* [4]
Coefficient estimates		
Estimation method	θ_0	θ_1
Dynamic OLS	2.16** (1.01)	0.86*** (0.24)
Johansen (1991) maximum likelihood	0.39 (1.21)	1.44***(0.29)

NOTE: The AEG and MZ_α statistics correspond to a one-sided (lower-tail) test of the null hypothesis that the 3-month Treasury bill rate and inflation rate are not cointegrated against the alternative hypothesis that the variables are cointegrated. The 10 percent, 5 percent, and 1 percent critical values for the AEG statistic are –3.07, –3.37, and –3.96; the 10 percent, 5 percent, and 1 percent critical values for the MZ_α statistic are –12.80, –15.84, and –22.84. The trace statistic corresponds to a one-sided (upper-tail) test of the null hypothesis that the 3-month Treasury bill rate and inflation rate are not cointegrated against the alternative hypothesis that the variables are cointegrated. The 10 percent, 5 percent, and 1 percent critical values for the trace statistic are 18.47, 20.66, and 24.18. The lag order for the regression model used to compute the test statistic is reported in brackets. *, **, and *** indicate significance at the 10 percent, 5 percent, and 1 percent levels. Standard errors are reported in parentheses.

nal interest rates argues for cointegration analysis of those variables to ascertain the EPRR's integration properties. When we prespecify a (1,–1)′ cointegrating vector and apply unit root tests to the EPRR, we reject the unit root null at the 5 percent level using the ADF statistic and at the 1 percent level using the MZ_α statistic. The U.S. EPRR appears to be stationary.

To test the null hypothesis of no cointegration without prespecifying a cointegrating vector, Table 3 reports the AEG statistic, MZ_α statistic from Perron and Rodriguez (2001), and trace statistic from Johansen (1991). The AEG and trace statistics reject the null hypothesis of no cointegration at the 10 percent level, and the MZ_α statistic rejects the null at the 5 percent level. Table 3 also reports estimates of the cointegrating coefficients, θ_0 and θ_1. Neither the dynamic ordinary least squares (OLS) nor Johansen (1991) estimates of θ_1 are significantly different from unity, indicating a stationary U.S. EPRR. The cointegrating vector is not estimated precisely enough to determine whether there is a tax effect.

Tables 2 and 3 provide evidence that the U.S. EPRR is stationary, although some of the rejections are marginal. Unit root and cointegration test results, however, are sensitive to the test procedure and sample period. Studies such as Mishkin (1992), Wallace and Warner (1993), and Crowder and Hoffman (1996) find evidence of a stationary U.S. EPRR, but Koustas and Serletis (1999) and Rapach and Weber (2004) generally do not. In contrast, per capita consumption growth is clearly stationary, as the ADF and MZ_α statistics in Table 2 both strongly reject the unit root null hypothesis for this variable. The fact that integration tests give mixed results for the EPRR's stationarity and clear-cut results for consumption growth highlights differences in the persistence properties of the two variables.

Confidence Intervals for the Sum of the Autoregressive Coefficients

The sum of the AR coefficients, ρ, in the AR representation of $i_t - \pi_{t+1}$ equals unity for an $I(1)$ process, while $\rho < 1$ for an $I(0)$ process. It is inherently difficult, however, to distinguish an $I(1)$ process from a highly persistent $I(0)$ process, as the two types of processes can be observationally equivalent (Blough, 1992; Faust, 1996).[12] To ana-

[12] In line with this, Crowder and Hoffman (1996) emphasize that impulse response analysis indicates that shocks have very persistent effects on the EPRR, although the U.S. EPRR appears to be $I(0)$.

lyze the theoretical implications of the time-series properties of the real interest rate, however, we want to determine a range of values for ρ that are consistent with the data, not only whether ρ is less than or equal to 1. That is, a series with a ρ value of 0.95 is highly persistent, even if it does not contain a unit root per se, and it is much more persistent than a series with a ρ value of, say, 0.4.

To calculate the degree of persistence in the data—rather than simply trying to determine if the series is $I(0)$ or $I(1)$—Rapach and Wohar (2004) compute 95 percent confidence intervals for ρ using the Hansen (1999) grid-bootstrap and Romano and Wolf (2001) subsampling procedures.[13] Using quarterly nominal long-term government bond yield and CPI inflation rate data for 13 industrialized countries for 1960-68, Rapach and Wohar (2004) report that the lower bounds of the 95 percent confidence interval for ρ for the tax-adjusted EPRR are often greater than 0.90, while the upper bounds are almost all greater than unity. Similarly, Karanasos, Sekioua, and Zeng (2006) use a long span of monthly U.S. long-term government bond yield and CPI inflation data for 1876-2000 to compute a 95 percent confidence interval for the EPRR's ρ. Their computed interval, (0.97, 0.99), indicates that the U.S. EPRR is a highly persistent or near-unit-root process, even if it does not actually contain a unit root.

With the same U.S. data underlying the results in Tables 2 and 3, we use the Hansen (1999) grid-bootstrap and Romano and Wolf (2001) subsampling procedures to compute a 95 percent confidence interval for ρ in the $i_t - \pi_{t+1}$ process. The grid-bootstrap and subsampling confidence intervals are (0.77, 0.97) and (0.71, 0.97), and the upper bounds are consistent with a highly persistent process. In contrast, the grid-bootstrap and subsampling 95 percent confidence intervals or ρ for per capita consumption growth are (0.34, 0.70) and (0.37, 0.64). The upper bounds of the confidence intervals for ρ for consumption growth are less than the lower bounds of the confidence intervals for ρ for the EPRR. This is another way to characterize the mismatch in the persistence properties of the EPRR and consumption growth.

Testing for Fractional Integration

Unit root and cointegration tests are designed to ascertain whether a series is $I(0)$ or $I(1)$, and the $I(0)/I(1)$ distinction implicitly restricts—perhaps inappropriately—the types of dynamic processes allowed. In response, some researchers test for fractional integration (Granger, 1980; Granger and Joyeux, 1980; Hosking, 1981) in the EARR and EPRR. A fractionally integrated series is denoted by $I(d)$, $0 \le d \le 1$. When $d = 0$, the series is $I(0)$, and shocks die out at a geometric rate; when $d = 1$, the series is $I(1)$, and shocks have permanent effects or "infinite memory." An intermediate case occurs when $0 < d < 1$: The series is mean-reverting, as in the $I(0)$ case, but shocks now die out at a much slower hyperbolic (rather than geometric) rate. Series in which $0 < d < 1$ exhibit "long memory," mean-reverting behavior, and can be substantially more persistent than even a highly persistent $I(0)$ series.

A number of studies, including Lai (1997), Tsay (2000), Karanasos, Sekioua, and Zeng (2006), Sun and Phillips (2004), and Pipatchaipoom and Smallwood (2008), test for fractional integration in the U.S. EPRR or EARR. Using U.S. postwar monthly or quarterly U.S. data, Lai (1997), Tsay (2000), and Pipatchaipoom and Smallwood (2008) all present evidence of long-memory, mean-reverting behavior, as estimates of d for the U.S. EPRR or EARR typically range from 0.7 to 0.8 and are significantly above 0 and below 1. Using a long span of annual U.S. data (1876-2000), Karanasos, Sekioua, and Zeng (2006) similarly find evidence of long-memory, mean-reverting behavior in the EPRR. Sun and Phillips (2004) develop a new bivariate econometric procedure that estimates the EARR's d parameter in the 0.75 to 1.0 range for quarterly postwar U.S. data.

Overall, fractional integration tests indicate that the U.S. EPRR and EARR do not contain a

[13] Andrews and Chen (1994) argue that the sum of the AR coefficients, ρ, characterizes the persistence in a series, as it is related to the cumulative impulse response function and the spectrum at zero frequency. While conventional asymptotic or bootstrap confidence intervals do not generate valid confidence intervals for nearly integrated processes (Basawa et al., 1991), Hansen (1999) and Romano and Wolf (2001) show that their procedures do generate confidence intervals for ρ with correct first-order asymptotic coverage. Mikusheva (2007) shows, however, that while the Hansen (1999) grid-bootstrap procedure has correct asymptotical coverage, the Romano and Wolf (2001) subsampling procedure does not.

unit root per se but are mean-reverting and very persistent. We confirm this by estimating d for the EPRR using our sample of U.S. data for 1953:Q1–2007:Q2 with the Shimotsu (2008) semiparametric two-step feasible exact local Whittle estimator that allows for an unknown mean in the series. This estimator refines the Shimotsu and Phillips (2005) exact local Whittle estimator, and these authors show that such local Whittle estimators of d have good properties in Monte Carlo experiments. The estimate of d for the EPRR is 0.71, with a 95 percent confidence interval of (0.51, 0.90), so we can reject the hypothesis that $d = 0$ or $d = 1$. This evidence of long-memory, mean-reverting behavior is consistent with the results from the literature discussed previously. The estimate of d for per capita consumption growth is 0.15 with a standard error of 0.10, so we cannot reject the hypothesis that $d = 0$ at conventional significance levels. This is another manifestation of the discrepancy in persistence between the real interest rate and consumption growth.

Testing for Threshold Dynamics and Nonlinear Cointegration

The empirical literature on the real interest rate typically uses models that assume both the cointegrating relationship and short-run dynamics to be linear.[14] Recently, researchers have begun to relax these linearity assumptions in favor of nonlinear cointegration or threshold dynamics, which allow for the cointegrating relationship or mean reversion to depend on the current values of the variables. For example, a threshold model might permit the EPRR to be approximately a random walk within ±2 percent of some long-run equilibrium value but to revert strongly to the ±2 percent bands when it wanders outside the bands.[15]

Million (2004) presents evidence that the U.S. EPRR adjusts in a nonlinear fashion to a long-run equilibrium level using a logistic smooth transition autoregressive (LSTAR) model and monthly U.S. 3-month Treasury bill rate and CPI inflation rate data for 1951-99. The Lagrange multiplier test of Luukkonen, Saikkonen, and Teräsvirta (1988) rejects the null hypothesis of a linear dynamic adjustment process, and there is evidence of stronger (weaker) mean reversion in the EPRR for values of the EPRR below (above) a threshold level of 2.2 percent. Million (2004) notes that the weak mean reversion in the upper regime is consistent with the fact that the U.S. real interest rate was persistently high during much of the 1980s, and he observes that the Federal Reserve's priority on fighting inflation, following the stagflation of the 1970s, could explain this period of high real rates. In a vein similar to that of Million, Koustas and Lamarche (2008) estimate three-regime self-exciting threshold autoregressive (SETAR) models to characterize the monetary policy strategy of "opportunistic disinflation" (Blinder, 1994; Orphanides and Wilcox, 2002). Based on the nonlinear unit root test of Bec, Salem, and Carassco (2004) and Hansen (1996, 1997) linearity tests, Koustas and Lamarche (2008) conclude that the EPRR can be suitably modeled as a three-regime SETAR process in Canada, France, and Italy over the postwar period.[16]

Christopoulos and León-Ledesma (2007) examine quarterly U.S. 3-month Treasury bill rate and CPI inflation rate data for 1960-2004, permitting the cointegrating relationship itself to be nonlinear. More precisely, they allow the cointegrating coefficient (θ_1) to vary with the inflation rate by estimating logistic and smooth exponential transition regression (LSTR and ESTR) models. Christopoulos and León-Ledesma (2007) find significant evidence of nonlinear cointegration between the nominal interest rate and inflation rate using the Choi and Saikkonen (2005) test. Using estimation techniques from Saikkonen and Choi (2004), the authors conclude

[14] Studies that allow for fractional integration or structural breaks also relax some linearity assumptions but in a different way than those reviewed in this subsection.

[15] The purchasing power parity literature often uses these threshold models (Sarno and Taylor, 2002).

[16] Maki (2003) uses the Breitung (2002) nonparametric procedure that allows for nonlinear adjustment dynamics to test for cointegration between the Japanese nominal interest rate and CPI inflation rate for 1972:01–2000:12. While Maki (2003) finds significant evidence of cointegration between the nominal interest rate and inflation rate using the Breitung (2002) test, he does not estimate the cointegrating vector, so it is not clear that the long-run equilibrium relationship is consistent with a stationary EPRR.

that the ESTR model fits best over the full sample (1960:Q1– 2004:Q4) and the first subsample (1960:Q1–1978:Q1), whereas the LSTR model fits best over the second subsample (1979:Q1–2004:Q4). The estimated ESTR model for 1960:Q1–1978:Q1 is not consistent with a stationary real EPRR for any inflation rate, and the estimated LSTR model for 1979:Q1–2004:Q4 is consistent with a stationary EPRR only when the inflation rate moves above approximately 3 percent.

In summary, recently developed econometric procedures provide some evidence of threshold behavior or nonlinear cointegration in the EPRR in certain industrialized countries. In some cases, the threshold models accord well with our intuition about changes in central bank policies. Although evidence of threshold behavior in real interest rates is potentially interesting, the models do not obviate the persistence in real interest rates, as there are still regimes where the real interest rate behaves very much like a unit root process.

TESTING FOR REGIME SWITCHING AND STRUCTURL BREAKS IN REAL INTEREST RATES

Building on the work of Huizinga and Mishkin (1986), another strand of the empirical literature tests for structural breaks in real interest rates. Accounting for such breaks can substantially reduce the persistence within the regimes defined by those breaks (Perron, 1989). Similarly, failing to account for structural breaks can produce spurious evidence of fractional integration (Jouini and Nouira, 2006).

Using quarterly U.S. 3-month Treasury bill rate and CPI inflation rate data for 1961-86, Garcia and Perron (1996) use Hamilton's (1989) Markov-switching approach to test for regime shifts in the U.S. EPRR. Specifically, they allow the unconditional mean of an AR(2) process to follow a three-state Markov process. The three estimated states correspond to high, middle, and low regimes with means of approximately 5.5 percent, 1.4 percent, and –1.8 percent, respectively. The filtered probability estimates show that the EPRR was likely in the middle regime from 1961-73, the low regime from 1973-81, and the high regime from 1981-86. There is very little persistence within each regime, as the estimated AR coefficients (ρ_1 and ρ_2 in equation (A1)) are near 0 within regimes. Overall, Garcia and Perron (1996) argue that the U.S. real interest rate occasionally experiences sizable shifts in its mean value, while the real interest rate is close to constant within the regimes.

Applications of Markov-switching models typically assume that the model is ergodic, so the current state will eventually cycle back to any possible state. Structural breaks have some similar properties to Markov-switching regimes, but they are not ergodic—they do not necessarily tend to revert to previous conditions. Because real interest rates in Garcia and Perron (1996) exhibit no obvious tendency to return to previous states, structural breaks might be considered more appropriate for modeling real interest rate changes than Markov switching. Bai and Perron (1998) develop a powerful methodology for testing for multiple structural breaks in a regression model, and Caporale and Grier (2000) and Bai and Perron (2003) apply this methodology to the mean of the U.S. EPRR. Both studies use quarterly U.S. short-term nominal interest rate and CPI inflation rate data for 1961-86, and the estimated break dates are very similar: 1967:Q1, 1972:Q4, and 1980:Q2 in Caporale and Grier (2000) and 1966:Q4, 1972:Q3, and 1980:Q3 in Bai and Perron (2003). The breaks correspond to a decrease in the mean EPRR in 1966/1967, a further decrease in 1972, and a sharp increase in 1980. Caporale and Grier argue that changes in political regimes—party control of the presidency and Senate—produce these regime changes.

Rapach and Wohar (2005) extend the work of Caporale and Grier (2000) and Bai and Perron (2003) by applying the Bai and Perron (1998) methodology to the EPRR in 13 industrialized countries using tax-adjusted nominal long-term government bond yield and CPI inflation rate data for 1960-98. They find significant evidence of structural breaks in the mean of the EPRR in each of the 13 countries. Rapach and Wohar (2005) also find that breaks in the mean inflation rate often coincide with breaks in the mean EPRR for each country's data. Furthermore, increases (decreases) in the mean inflation rate are almost always associated with decreases (increases) in the mean EPRR.

Table 4

Bai and Perron (1988) Test Statistics and Estimation Results for the U.S. ex post Real Interest Rate (1953:Q1–2007:Q2)

Test statistic		Regime	Estimated ex post real interest rate mean
UD_{max}	14.84***	1953:Q1–1972:Q3 [1969:Q2, 1973:Q4]	1.22*** (0.17)
WD_{max} (5%)	27.06**	1972:Q4–1980:Q3 [1979:Q1, 1980:Q4]	–0.55 (0.38)
$F(1\|0)$	12.92***	1980:Q4–1989:Q3 [1984:Q3–1994:Q2]	4.58*** (0.71)
$F(2\|1)$	17.89***	1989:Q4–2007:Q2	1.82*** (0.52)
$F(3\|2)$	17.89***		
$F(4\|3)$	10.37*		
$F(5\|4)$	10.37		

NOTE: *, **, and *** indicate significance at the 10 percent, 5 percent, and 1 percent levels. The bracketed dates in the Regime column denote a 90 percent confidence interval for the end of the regime. Numbers in parentheses in the last column denote standard errors for the estimated mean.

This finding is consistent with the hypothesis that monetary easing increases inflation and generates a persistent decline in the real interest rate.

In a comment on Rapach and Wohar (2005), Caporale and Grier (2005) examine whether political regime changes affect the mean U.S. EPRR, after controlling for the effects of regime changes in the inflation rate. Caporale and Grier (2005) find that political regime changes associated with changes in the party of the president or control of Congress do not affect the mean EPRR after controlling for inflation. However, the appointments of Federal Reserve Chairmen Paul Volcker in 1979 and Alan Greenspan in 1987 are associated with shifts in the mean EPRR even after controlling for changes in the mean inflation rate.

The previous papers test for structural breaks under the assumption of stationary within-regime behavior. In the spirit of Perron (1989), a number of studies test whether the real interest rate is $I(0)$ after allowing for deterministic shifts in the mean real rate. Extending the methodology of Perron and Vogelsang (1992), Clemente, Montañés, and Reyes (1998) test the unit root null hypothesis for the U.K. and U.S. EPRR using quarterly long-term government bond yield and CPI inflation rate data for 1980-95, allowing for two breaks in the mean of the EPRR. They find that the EPRR in the United Kingdom and United States is an $I(0)$ process around an unconditional mean with two breaks. Using monthly U.S. 1-year Treasury bill rate data for 1978-2002 and expected inflation data from the University of Michigan's *Survey of Consumers*, Lai (2004) finds that the EARR is an $I(0)$ process with a shift in its unconditional mean in the early 1980s. Lai (2008) extends Lai (2004) by allowing for a mean shift in quarterly real interest rates for eight industrialized countries and eight developing countries and finds widespread support for a stationary EPRR after allowing for a break in the unconditional mean.

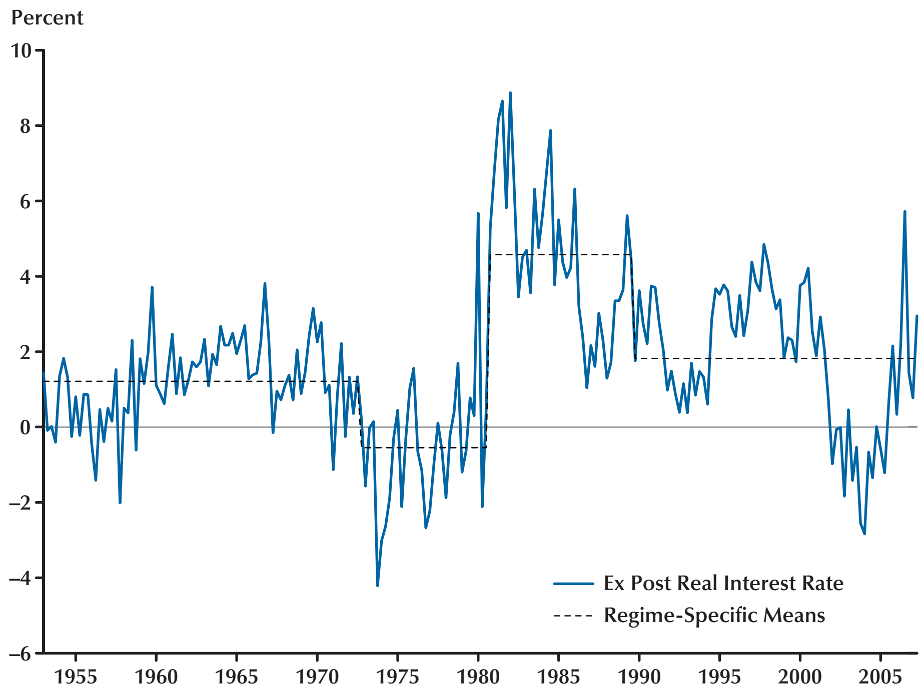
To further illustrate the prevalence of structural breaks, we use the Bai and Perron (1998) methodology to test for such instability in the unconditional mean of the U.S. EPRR for 1953:Q1–2007:Q2.[17] Table 4 reports the results. The procedure finds three changes in the mean that occur at 1972:Q3, 1980:Q3, and 1989:Q3 and are similar to those previously identified for the United States.[18] The breaks are associated with substan-

[17] We focus on the Bai and Perron (1998) methodology in analyzing mean real interest rate shifts in updated U.S. data. It would be interesting in future research to consider regime-switching models and recently developed structural break tests such as described by Elliott and Müller (2006).

[18] Rapach and Wohar (2005) discuss how the statistics reported in Table 4 imply that there are three significant breaks in the unconditional mean.

Figure 2

U.S. Ex Post Real Interest Rate and Regime-Specific Means, 1953:Q1–2007:Q2



NOTE: The figure plots the U.S. ex post real interest rate and means for the regimes defined by the structural breaks estimated using the Bai and Perron (1998) methodology.

tial changes in the average annualized real interest rate in the different regimes. The average real rate is 1.22 percent for 1953:Q1–1972:Q3, is not significantly different from zero for 1972:Q4–1980:Q3, increases to 4.58 percent for 1980:Q4–1989:Q3, and falls to 1.82 percent for 1989:Q4–2007:Q2. Figure 2 depicts the EPRR and the mean for each of the four regimes defined by the three breaks.[19] In contrast to this evidence for breaks in the real rate, the Bai and Perron (1998) methodology fails to discover significant evidence of structural breaks in the mean of per capita consumption growth. (We omit complete results for brevity.)

In interpreting structural break results, we emphasize that such breaks only reduce within-regime or local persistence in real interest rates. The existence of breaks still implies a high degree of global persistence, and the breaks themselves require an economic explanation.

THEORETICAL IMPLICATIONS AND A MONETARY EXPLANATION OF PERSISTENCE

This section considers what types of shocks are most likely to produce the persistence in the U.S. real interest rate. The empirical literature devotes relatively little attention to this important issue. We argue that monetary shocks likely drive the persistence in the U.S. real interest rate.

[19] The test results of Bai and Perron (1998) for structural breaks in the mean EPRR do not appear sensitive to whether the tax-adjusted or tax-unadjusted EPRR is used (Rapach and Wohar, 2005). Neither do estimates of the sum of the AR coefficients nor tests for fractional integration hinge critically on whether the EPRR is tax adjusted.

Before discussing potential sources of real interest rate persistence, we briefly make the case that the U.S. real interest rate is ultimately mean-reverting. As we emphasize, unit root and cointegration tests have difficulty distinguishing unit root processes from persistent but stationary alternatives. Nevertheless, unit root and cointegration tests with good size and power, applied to updated data, provide evidence that the U.S. real interest rate is an $I(0)$—and thus mean-reverting—process (see Table 2).[20] Tests for fractional integration nest the $I(0)/I(1)$ alternatives, and they concur that the U.S. real interest rate is a mean-reverting process. Using an updated sample, we confirm the findings of Lai (1997), Tsay (2000), Pipatchaipoom and Smallwood (2008), and Karanasos, Sekioua, and Zeng (2006) that demonstrate long-memory, mean-reverting behavior in the U.S. real interest rate. Our updated sample also provides evidence of structural breaks in the U.S. real interest rate. Curiously, the regime-specific mean breaks for the EPRR largely cancel each other in the long run (see Table 4): The estimated mean real rate in 2007 is close to that estimated for 1953.[21] We speculate that although structural breaks appear to describe the data better than a constant, linear data generating process, these breaks appear to exhibit a certain type of mean-reverting behavior. With sufficient data—much more than we have now—one could presumably model this mean-reversion in regimes.

These facts lead us to tentatively claim that the U.S. real interest rate is best viewed as a very persistent but ultimately mean-reverting process. We emphasize the tentative nature of this claim, and we consider careful econometric testing of this proposition to be an important area for future research. Even if real interest rates ultimately mean-revert, they are clearly very persistent.

Recall the underlying motivation for learning about real interest rate persistence: In a simple endowment economy, the real interest rate should have the same persistence properties as consumption growth. In fact, however, real rates are much more persistent than consumption growth. Permanent technology growth shocks can create a nonstationary real rate but affect consumption growth in the same way, so they cannot account for the mismatch in persistence. More complex equilibrium growth models potentially explain this persistence mismatch through changing fiscal and monetary policy, as well as transient technology growth shocks. We consider fiscal, monetary, and transient technology shocks as potential causes of persistent fluctuations in the U.S. real interest rate.

Figures 1 and 2 reveal two episodes of pronounced and prolonged changes in the U.S. EPRR: the protracted decrease in the EPRR in the 1970s and subsequent sharp increase in the 1980s. Fiscal shocks appear to be an unlikely explanation for the large decline in real rates from 1972-79. The U.S. did not undertake the sort of contractionary fiscal policy that would be necessary for such a fall in real rates.[22] In fact, fiscal policy in the 1970s largely tended toward modest deficits. Given the substantial budget deficits beginning in 1981, expansionary fiscal shocks are a more plausible candidate for the increase in real rates at this time.

Monetary shocks appear to fit well with the overall pattern in the real interest rate, including the multiyear decline in the real rate during the 1970s, the very sharp 1980 increase, and subsequent gradual decline during the "Great Disinflation." One interpretation of the "Great Inflation" that began in the late 1960s and lasted throughout the 1970s is that the Federal Reserve pursued an expansionary monetary policy—either inadvertently or to reduce the unemployment rate to unsustainable levels—and this persistently reduced the real interest rate (Delong, 1997; Barsky and Kilian, 2002; Meltzer, 2005; Romer, 2005). After Paul Volcker's appointment as Chairman, the Federal Reserve sharply raised short-term nominal interest rates to reduce inflation from its early 1980 peak of nearly 12 percent, and this

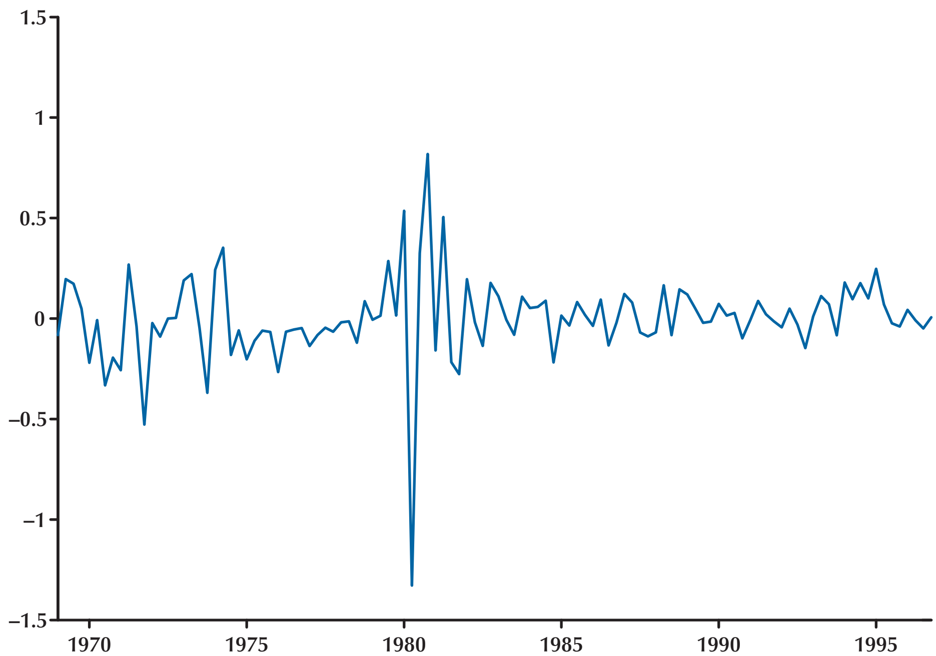
[20] Recall, however, that unit root and cointegration tests are sensitive to the particular sample used.

[21] One might wonder if the observed mean-reversion in structural breaks contradicts our contention that the breaks should not be modeled as a Markov process because they are not ergodic. We do not think, however, that observing one state twice and two states once provides sufficient information for a Markov process.

[22] The recent analyses by Romer and Romer (2008) and Ramey (2008) indicate that the U.S. economy did not experience sizable contractionary fiscal policy shocks during the 1970s.

Figure 3

Romer and Romer (2004) Measure of Monetary Policy Shocks, 1969:Q1–1996:Q4



NOTE: A positive (negative) value corresponds to a contractionary (expansionary) monetary policy shock.

produced a sharp and prolonged increase in the real interest rate. The structural breaks manifest these pronounced swings: The mean EPRR falls from 1.22 percent in 1972:Q3 to essentially zero and then rises to 4.58 percent beginning in 1980:Q4 (see Table 4). Furthermore, Rapach and Wohar (2005) report evidence of breaks in the mean U.S. inflation rate in 1973:Q1 and 1982:Q1 that increase and decrease the average inflation rate. The timing and direction of the breaks are consistent with a monetary explanation that also accounts for the mismatch in persistence between the real interest rate and consumption growth. In each case, negative (positive) breaks to the real rate of interest coincide with positive (negative) breaks in the mean rate of inflation. The data are in line with the hypothesis that central banks change monetary policy and inflation through persistent effects on the real rate of interest.

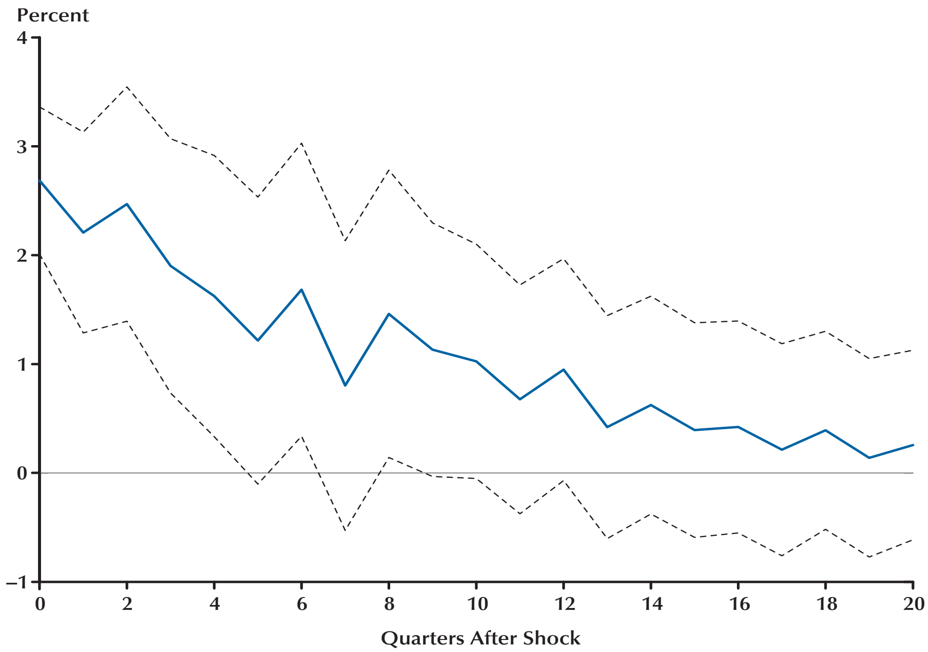
Turning to technology shocks, the paucity of independent data on technology shocks makes it difficult to correlate such changes with real interest rates. In addition, researchers have traditionally viewed technology growth as reasonably stable. One might think that other sorts of supply shocks, such as oil price increases, might influence the real rate, and they surely do to some degree; Barro and Sala-i-Martin (1990) and Caporale and Grier (2000), for example, consider this possibility. It is unlikely, however, that oil price shocks alone can account for the pronounced swings in the U.S. real interest rate: Why would rising oil prices in 1973 reduce the real interest rates but rising oil prices in 1979 dramatically raise the real rate?[23]

While we interpret the timing of major swings in the U.S. real rate to strongly suggest a monetary explanation, we ultimately need to estimate

[23] Furthermore, Barsky and Kilian (2002) argue that the timing of increases in U.S. inflation in the early 1970s is more consistent with a monetary rather than an oil price shock explanation.

Figure 4

U.S. Ex Post Real Interest Rate Response to a Contractionary Romer and Romer (2004) Monetary Policy Shock



NOTE: The response is based on an autoregressive distributed lag model estimated for 1969:Q1–1996:Q4. Dashed lines delineate two-standard-error bands. The response is to a shock of size 0.5.

structural models to analyze the relative importance of various shocks. Galí (1992) is one of the few studies providing evidence on the economic sources of real interest rate persistence. His SVAR model finds that an expansionary money supply shock leads to a very persistent decline in the real interest rate, and money supply shocks account for nearly 90 percent of the variance in the real rate at the one-quarter horizon and still account for around 60 percent of the variance at the 20-quarter horizon. Galí's (1992) evidence is consistent with our monetary explanation of real interest rate persistence.[24]

We present tentative additional evidence in support of a monetary explanation of real interest rate persistence based on the new measure of monetary shocks developed by Romer and Romer (2004). They cull through quantitative and narrative Federal Reserve records to compute a monetary policy shock series for 1969-96 that is independent of systematic responses to anticipated economic conditions. Figure 3 plots the Romer and Romer (2004) monetary policy shocks series, where expansionary (i.e., negative) shocks in the late 1960s and early 1970s and large contractionary (i.e., positive) shocks in the late 1970s and early 1980s appear to match well with the decline in the U.S. real interest rate in the 1970s and subsequent sharp increase around 1980.

Romer and Romer (2004) estimate autoregressive distributed lag (ARDL) models to examine the effects of a monetary policy shock on real output and the price level. They find that a contractionary shock creates persistent and sizable

[24] King and Watson (1997) and Rapach (2003) use SVAR frameworks to estimate the long-run effects of exogenous changes in inflation on the real interest rate. Both studies find evidence that an exogenous increase in the steady-state inflation rate decreases the steady-state real interest rate.

declines in both real output and the price level. In similar fashion, we estimate an ARDL model via OLS to measure the effects of a monetary policy shock on the real interest rate. The ARDL model takes the form,

$$r_t^{ep} = a_0 + \sum_{j=1}^{8} a_j r_{t-j}^{ep} + \sum_{j=0}^{8} b_j S_{t-j} + u_t, \tag{8}$$

where r_t^{ep} is the EPRR and S_t is the Romer and Romer measure of monetary policy shocks.

Figure 4 illustrates the response of the EPRR to a monetary policy shock of size 0.5, which is comparable to some of the contractionary shocks experienced in the late 1970s and early 1980s (see Figure 3). Romer and Romer's (2004) Monte Carlo methods provide the two-standard-error bands. A contractionary monetary policy shock produces a statistically and economically significant increase in the U.S. EPRR, which remains statistically significant after approximately two years. Note that the response in Figure 4 is nearly identical to the response of r_t^{ep} to a shock to S_t obtained from a bivariate VAR(8) model that orders S_t first in a Cholesky decomposition. Together, Figures 3 and 4 show that expansionary (contractionary) monetary policy shocks can account for the pronounced and prolonged decrease (increase) in the U.S. real interest rate in the 1970s (early 1980s). We emphasize that this evidence is suggestive. Of course, structural identification is a thorny issue, and more research is needed to determine the veracity of the monetary explanation for U.S. real interest rate persistence.

CONCLUSION

Rose's (1988) seminal study spurred a sizable empirical literature that examines the time-series properties of real interest rates. Our survey details the evidence that real interest rates are highly persistent. This persistence manifests itself in the following ways:

- Under the assumption of a constant data generating process, many studies indicate that real interest rates contain a unit root. While econometric problems prevent a dispositive resolution of this question, real interest rates display behavior that is very persistent, close to a unit root.
- Estimated 95 percent confidence intervals for the sum of the AR coefficients from the literature have upper bounds that are greater than or very near unity.
- Real interest rates appear to display long-memory behavior; shocks are very long-lived, but the real interest rate is estimated to be ultimately mean-reverting.
- Studies allowing for nonlinear dynamics in real interest rates identify regimes where the real interest behaves like a unit root process.
- Structural breaks in unconditional means characterize real interest rates. Although the breaks reduce within-regime persistence, the real interest rate remains highly persistent because the regimes have different means.

Although researchers have used a variety of econometric models to analyze the time-series properties of real interest rates, relatively little work has been done to discriminate among these sundry models. Model selection could tell us, for example, whether we should think of persistent changes in real interest rates in terms of changes in the steady-state real rate—which are consistent with unit root behavior—or long-lived shocks that eventually decay to a stable steady-state real rate—which are consistent with mean-reverting behavior. While model selection raises challenging econometric (and philosophical) issues, out-of-sample forecasting exercises and analysis of posterior model probabilities in a Bayesian context might identify the best way to model real interest rate persistence.

Finally, structural analysis is necessary to identify the sources of the persistence in real interest rates. Theoretical models suggest that a variety of shocks can induce real rate persistence, including preference, technology growth, fiscal, and monetary shocks. We suggest a tentative monetary explanation of U.S. real interest rate persistence based on timing, lack of persistence in consumption growth, and large and persistent real interest rate responses to a Romer and Romer (2004) monetary policy shock. The literature would greatly benefit from further analysis of the relative importance of different types of shocks in explaining real interest rate persistence.

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APPENDIX

Unit Root and Cointegration Tests

This appendix briefly describes the basic framework for unit root and cointegration testing; Hamilton (1994) details the subject. Following Dickey and Fuller (1979) and Said and Dickey (1984), unit root tests are typically based on the autoregressive (AR) representation of a time series, which can be written as follows:

(A1)
$$y_t - \mu = \rho_1\left(y_{t-1} - \mu\right) + \ldots + \rho_k\left(y_{t-k} - \mu\right) + e_t,$$

where e_t is a white noise disturbance term. When the sum of the AR coefficients in equation (A1),

$$\rho = \sum_{j=1}^{k} \rho_j,$$

equals 1, shocks to y_t persist forever—y_t has a unit root and thus has no tendency to revert to an unconditional mean. Testing the null hypothesis that $y_t \sim I(1)$ against the alternative hypothesis that $y_t \sim I(0)$ is equivalent to testing

$$\sum_{j=1}^{k} \rho_j = 1 \text{ versus } \sum_{j=1}^{k} \rho_j < 1.$$

Researchers usually ignore the possibility that $\rho > 1$, since this would imply an explosive process, which we do not observe in the data. The *t* statistic on γ in the following augmented Dickey-Fuller (ADF) regression provides a convenient test statistic for the unit root null hypothesis:

(A2)
$$\Delta y_t = \delta + \gamma y_{t-1} + \tilde{\rho}_1 \Delta y_{t-1} + \ldots + \tilde{\rho}_{k-1} \Delta y_{t-(k-1)} + e_t,$$

where $\delta = \mu(1-\rho)$, $\gamma = -(1-\rho)$, and

$$\tilde{\rho}_i = -\sum_{j=i+1}^{k} \rho_j.$$ [25]

Under the null hypothesis that $y_t \sim I(1)$, $\gamma = 0$, while $\gamma < 0$ under the alternative hypothesis that $y_t \sim I(0)$. The *t* statistic on γ in equation (A2) has a nonstandard distribution, necessitating simulation methods to obtain critical values.

Cointegration tests are closely related to unit root tests in that they ask whether any linear combination of some set of $I(1)$ processes (say, y_t and x_t) are stationary or cointegrated. The popular, residual-based augmented Engle and Granger (1987, AEG) procedure uses the following ordinary least squares (OLS) regression as a first step in testing the null hypothesis of no cointegration:

(A3)
$$y_t = \theta_0 + \theta_1 x_t + u_t.$$

The cointegrating vector, which defines the stable long-run relationship between y_t and x_t (if it exists), is given by $(1,-\theta_1)'$. One then runs an ADF-type unit root test—with no constant—on the regression residuals, $\hat{u}_t = y_t - (\hat{\theta}_0 + \hat{\theta}_1 x_t)$, where $\hat{\theta}_0$ and $\hat{\theta}_1$ are the OLS estimates of θ_0 and θ_1. The AEG test statistic—the ADF test statistic from the residual regression—also has a nonstandard asymptotic distribution, which requires simulated critical values. When y_t and x_t are cointegrated, $\hat{\theta}_0$ and $\hat{\theta}_1$ are superconsistent, converging to their probability limits faster than the usual rate of

$$1/\sqrt{T}.$$

[25] The unit root tests developed by Phillips and Perron (1988) are closely related to ADF tests and are frequently used in the literature. We refer to both ADF and Phillips and Perron (1988) tests simply as ADF tests in our discussion of the empirical literature in the text.

Endogeneity bias, however, renders conventional OLS standard errors incorrect. When y_t and x_t are cointegrated, fully modified OLS (FM-OLS; Phillips and Hansen, 1990) and dynamic OLS (DOLS; Saikkonen, 1991; Stock and Watson, 1993) procedures efficiently estimate θ_0 and θ_1 with appropriate standard errors.

Johansen (1991) develops a cointegration test procedure based on the likelihood function of a system of equations that simultaneously tests the null hypothesis of no cointegration and consistently and efficiently estimates the cointegrating vector (if it exists). This system-based approach is also popular in applied research and is potentially more powerful than the single-equation–based AEG approach (Pesavento, 2004).

Unit root and cointegration tests have two significant problems. First, they have low power to reject the null if the true model is a highly persistent but stationary process (DeJong et al., 1992). Second, moving-average components in the underlying data-generating process complicate inference from unit root and cointegration tests. Schwert (1987, 1989) shows that ADF unit root tests can have substantial size distortions that lead to spurious rejections of the unit root null hypothesis in the presence of a significant moving-average component.[26] Lütkepohl and Saikkonen (1999) show that such size distortions can also affect cointegration tests. This is potentially relevant when analyzing the EPRR, as Perron and Ng (1996) and others show that inflation rates often have sizable moving-average components.[27]

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[26] There are two strategies for dealing with a significant moving-average component in the data-generating process when performing ADF unit root tests: (i) include a large number of lags when estimating (A2), as an autoregressive moving average process with finite-order lag polynomials can be expressed as an infinite-order AR process; (ii) include the moving-average component in the data generating process when simulating critical values.

[27] Perron (1994) observes that the inflation rate could exhibit a substantial moving-average component if the monetary authority offsets inflationary or disinflationary shocks away from a target price level path.

Drug Prices Under the Medicare Drug Discount Card Program

Emin M. Dinlersoz, Rubén Hernández-Murillo, Han Li, and Roger Sherman

In early 2004, the U.S. government initiated the Medicare Drug Discount Card Program (MDDCP), which allowed card subscribers to obtain discounts on prescription drugs. Pharmacy-level prices were posted on the program website weekly with the hope of promoting competition among card sponsors by facilitating consumer access to prices. A large panel of pharmacy-level price data collected from this website indicates that price dispersion across cards persisted throughout the program. Prices declined initially when consumers were choosing cards, but rose later when subscribers were restricted to commit to their card choices. In contrast, contemporaneous prices from online drug retailers, which were unrelated to the program, rose steadily over time, indicating that program prices evolved in a way different from the general evolution of prices outside the program. (JEL D43, D83, I11, I18, L11, L13, L50)

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On April 29, 2004, in conjunction with the Medicare Drug Discount Card Program and Transitional Assistance Program (MDDCP), the U.S. government activated a website to publicize prices offered by discount cards for more than 800 prescription drugs at individual pharmacy levels across all zip code areas in the United States. The MDDCP was initiated as a transition to the broader Medicare Part D prescription drug assistance program that took effect in January 2006, aiming to lower the cost of drugs and therapy for elderly and handicapped individuals covered by Medicare. The price information on the MDDCP website was updated on a weekly basis for the duration of the program. This mandatory release of prices continues under the Medicare Part D program, and it is unmatched in scale in the history of government policy on information transparency.

The MDDCP and its successor program, Medicare Part D, were intended to induce competition among drug card sponsors, largely through the extensive amount of price information that drug card sponsors were required to release on the website. The premise was that the ease of consumer search for prices in the program website would enable them to choose the lowest-price card sponsor, leading to intensified competition among sponsors. However, at the same time that the MDDCP generated price information with the intent to boost competition among drug cards, the program design also required subscribers to commit to a single drug card once they subscribed, rather than being allowed to switch cards at will. This institutional constraint on consumer switching, among other factors, could inhibit competi-

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tion by preventing consumers from switching to low-price providers. Thus, a major question is whether MDDCP competition among the drug card sponsors was indeed effective in lowering drug prices as intended by the program initiators.

This paper studies the dynamics of prices over the course of the program by using a large sample of prices collected from the MDDCP website for several weeks. The empirical analysis indicates that the program resulted in economically significant and persistent price dispersion across cards. More importantly, the evidence points to a nonmonotonic time path for prices. Drug prices declined in the early phases of the program when card subscription was still diffusing across consumers, and they rose later when new card subscriptions slowed and consumers could no longer switch cards, although the magnitudes of these shifting trends were not exceptionally large relative to the overall average of program prices. As a benchmark for comparison, contemporaneous prices unrelated to the program were collected from online drug retailers, and these prices exhibited a steady upward trend. In particular, when MDDCP prices declined, online prices rose, and when both sets of prices rose, the rise in MDDCP prices was actually greater than the rise in online prices. Thus, MDDCP prices evolved differently from the general evolution of drug prices outside the MDDCP, indicating that the time path of prices within the MDDCP cannot be explained simply by general trends in regular online drug prices.

The analysis of the evolution of prices under the MDDCP is relevant because of the potentially large welfare consequences of government policies aimed at increasing competition, which continue under Medicare Part D. At the time of the study, the population eligible for drug cards was around 7.5 million, and it has continued to grow since Medicare Part D took effect in 2006. The design of a viable prescription drug program for the elderly is still a major policy issue and the success of the ongoing Medicare Part D remains to be seen. Therefore, lessons learned from the MDDCP experience are valuable in assessing the success of government-sponsored competition and information dissemination about prices. By increasing program awareness, making the price information publicly available through its website, and helping eligible consumers choose cards, the program's goal was to increase competition among rival prescription drug suppliers and to help establish a market in which such increased competition persisted in the long run even after the broader Medicare Part D took effect. The belief, at least in part, was that effective consumer search for lower prices would discipline the pricing behavior of suppliers and lead to lower prices. Certain theories of consumer search (e.g., Stahl, 1989) suggest that easier consumer search can exert downward pressure on prices in a market. Related empirical work (notably, Brown and Goolsbee, 2002) also demonstrated that the diffusion of consumer search can be associated with lower prices, provided consumers were uninhibited in switching to lower-priced suppliers. Yet, the results here suggest that no systematic and economically significant decline in prices occurred when the overall price dynamics within the program are considered.

Official surveys and studies on the marketing performance of the MDDCP point to several reasons that appear to have prohibited effective search by consumers, which could have allowed the reallocation of consumers to lower-priced cards.[1] Efforts to inform potential beneficiaries about drug cards and the enrollment process were limited, particularly about the cost of cards and the extent of the discounts available. In addition, the program's website and the existing help lines were not particularly useful in guiding consumers to choose the card that was best for them. Thus, these studies do not suggest any strong indication that consumers were able to make highly informed and close to ideal choices. Consumers also were often confused about the abundance of cards and pharmacies from which to choose, which made it difficult to make the best choice. Furthermore, the low inherent propensity of Internet usage and searching capability among elderly individuals prevented effective usage of the website for consumer decisionmaking. All these impediments

[1] See, for instance, the 2005 U.S. Government Accountability Office study GAO-06-13R.

seem to have led to price dynamics in the program that did not coincide with the program's intended effects.

The rest of this paper is organized as follows. The next section provides background for the MDDCP. The third section presents some theoretical guidance, followed by a description of the data. Empirical analysis and results are then described before our concluding section.

THE MDDCP BACKGROUND

The design and the institutional environment of the MDDCP are crucial for understanding the functioning of the retail drug markets created by the program. The MDDCP allowed qualified drug card sponsors to make arrangements with drug manufacturers to obtain discounts and pass these discounts on to Medicare recipients. Eligible consumers could subscribe on a strictly voluntary basis to a card of their choice and obtain their prescriptions at a discount specified by the card sponsor. Prescriptions were available either from retail pharmacies or by mail from mail-order pharmacies that had arrangements with the card sponsor. An individual consumer subscribed to a card by paying a fixed annual fee (for at most two years), ranging between $0 and $30, and thereafter was entitled to receive that card's discounts. The consumer's problem consisted of two stages: first, choosing a drug card that provided the best discount on the bundle of drugs used by the consumer and second, choosing a retail (or mail-order) pharmacy that sold the drugs of interest.

Certain institutional aspects of the program were relevant for the dynamics of program prices. First, a card sponsor was not required to commit to a given level of discount on drugs over time. This flexibility in card prices left the door open for price fluctuations that could result from competition among cards, above and beyond general fluctuations in drug prices such as those related to changes in manufacturers' costs or changes in demand after the introduction of a generic. For a given card, there was also no prior commitment for prices to be the same across all pharmacies that offered discounts under the card.

Second, in addition to the usual consumer search and switching costs that contribute to price dispersion in drug retail markets (see Scott-Morton, 1997, and Sorensen, 2000, 2001), prohibitive consumer switching costs were inherent in the very design of the program.[2] Once enrolled in a card program, a consumer was not allowed to switch to another card, except in certain special cases, such as moving to a new location or a card sponsor exiting the market. This restriction on switching introduced additional friction and inertia into the market, which may have impeded reallocation of consumers to low-price card sponsors over time. The MDDCP had a nationally coordinated switching period between November 15 and December 31, 2004, during which consumers were allowed to review their card choices and change them if they wished to do so. After this period, a consumer who was already enrolled in a card was not allowed to switch to another card until the end of the program, subject to the exceptions mentioned. The prevention of switching after the switching period and the timing of the switching period could potentially lead to price dynamics driven by the card sponsors' incentives to charge lower prices in the early stages of the program to attract subscribers, and then to increase their prices once consumers were locked in to their card choices.

Third, the diffusion of card enrollment among eligible consumers was expected to be gradual, not instantaneous. Consumers had to evaluate card choices before making a decision. One of the main criticisms of the MDDCP was the complexity of the card choice process related to the abundance of alternative plans whose benefits were hard to assess. This criticism applies equally to Medicare Part D. Available evidence indicated that the diffusion of card enrollment was indeed gradual. According to enrollment data from the Center for Medicare and Medicaid Services (CMS), about 6.4 million beneficiaries were enrolled in the drug card program as of September 2005,

[2] Usual switching costs in the context of prescription drugs include consumer learning costs about the side effects of a new drug that can substitute for the consumer's existing drug and physicians' inertia in changing prescriptions because of rewards and loyalty programs offered by the manufacturer or the wholesaler of that drug.

toward the end of the program.[3] Roughly two-thirds of participants enrolled early in the program (May through July 2004). Enrollment was much faster between May and October 2004 and reached about 6 million participants (about 80 percent of the total Medicare population) around October 2004. It rose little thereafter, essentially staying level after January 2005, when the switching period ended.

Moreover, most consumers eligible for cards were 65 years or older, not a group of particularly Internet-savvy consumers. Shortly before the program took effect, Fox (2004) estimated that 22 percent of adults aged 65 and older had access to the Internet. Of this group, an estimated 66 percent used the Internet to locate health information, implying that only about 14 percent of the relevant population used the Internet for health information searches. Thus, the overall propensity to use the Internet as a price search tool was not impressively high in the eligible population.

Further evidence of consumers' enrollment and experience with the program comes from an October 2005 report on the progress of the MDDCP program prepared by Abt Associates, Inc., on request from the CMS.[4] Based on an extensive survey of card enrollees and non-enrollees, the report found that widespread awareness of the MDDCP was obtained within a few months of the program. Although a majority of respondents reported that they had more than enough information to make a choice among the cards, one quarter to half did not consider more than just one drug card. Some consumers simply took the first card available, whereas others were enrolled automatically. About 13 percent of survey participants obtained information from the Medicare website, either directly or with the help of a family member, friend, or counselor who accessed the website for them. Overall, the available evidence indicates that both the rate of learning about cards and the search rate for lower prices were rather low.

THEORETICAL CONSIDERATIONS

Consumer search is an important source of price dispersion in retail drug markets (e.g., Sorensen, 2000, 2001). Static models of search are abundant in the literature (see, e.g., Salop and Stiglitz, 1977, Reinganum, 1979, Burdett and Judd, 1983, and Stahl, 1989). For instance, Stahl (1989) shows that as the proportion of consumers who are fully informed of prices increases, average price falls monotonically. Price dispersion exhibits nonmonotonic behavior, initially increasing for low values of the informed proportion, but decreasing for higher values. Although comparative statics from this static model can be used, as in Brown and Goolsbee (2002), to draw some conclusions for a dynamic framework, the MDDCP's institutional environment introduces further considerations for firms' and consumers' behavior over time, which call for a dynamic framework.

Given the available evidence on intensity of search discussed in the previous section, it is hard to argue that consumer search worked as effectively as in ordinary, nonprogram retail prescription drug markets. Factors in the previous section suggest that there may have been little consumer search in the market created by MDDCP, and the abundance of choices and the complexities of the program design may have inhibited search. Another major constraint of the program is that it prevented consumers from using more than one card or from changing their card choices after they subscribed, with few exceptions. The prohibitive switching cost could have induced card sponsors to lower their prices in the early stages of the program to attract consumers who had not yet chosen a card. But as more and more consumers were locked in to their choices, card sponsors would have incentives to raise prices. After the switching period, prices may be expected to rise as sponsors take advantage of consumers' inability to change cards.

This nonmonotonic time path of prices indeed arises in certain models of dynamic price competition with consumer switching costs, such as those of Klemperer (1987) and Farrell and Shapiro (1988). The MDDCP had a lifetime of less than

[3] For more details of the enrollment patterns, see the 2005 U.S. Government Accountability Office study GAO-06-13R.

[4] See Hassol et al. (2005).

two years, and cards were differentiated in many dimensions beyond just price. These main features of the program are captured nicely by the model of Klemperer (1987), which presents a two-period differentiated-products duopoly in which consumers are partially locked in by switching costs that they face in the second period. Switching costs make demand more inelastic in the second period. Prices are lower in the first period as firms compete to build a customer base that is valuable later. However, prices may be higher in both periods than they would be in a market without switching costs.

Two main considerations under the MDDCP may make the price dynamics differ from those in Klemperer (1987). First, Klemperer (1987) assumes perfect consumer information about prices, whereas the evidence discussed above suggests that many card enrollees under the MDDCP chose their drugs with imperfect information about the cards' benefits and prices. Lack of perfect information about prices would not change the competition in the second period, because the constraints on switching would prevent consumers from abandoning their firms even if they were informed of a lower price at some point. However, the intensity of competition in the first period could change. Firms could take advantage of consumers' imperfect information and not lower their prices as much as they would in the case of perfect information. Obviously, a related issue is that each card sponsor itself probably did not have good information on the general pattern of card enrollment and on imperfections in consumers' information about cards. If card sponsors believed, at least initially until firm evidence on enrollment patterns emerged, that consumers would make informed decisions, they would have incentives to lower their prices.

Second, the MDDCP's allowance for a round of card switching in the middle of the program created incentives for a potential price war by card sponsors. One implication is that, in addition to lower prices at the early phases of the program, lower prices would be expected during the switching period compared with nonswitching periods. There is no artificially introduced "switching period" in Klemperer (1987).

Other considerations, however, could prevent this predicted nonmonotonic path for prices. Given the continuing nature of the prescription drug program with Part D, card sponsors who use bait-and-switch strategies could harm their reputations. Although the MDDCP itself lasted only two years, many card sponsors continued to participate in Medicare Part D when it started in January 2006, so sponsors faced the possibility of alienating consumers because of bait-and-switch price strategies. One of the program's goals, as stated in the Medicare program–related website, was to prevent bait-and-switch behavior. However, the program did not spell out any strict guidelines as to what exactly constitutes bait-and-switch and how it would be prevented.

The discussion so far suggests that the level of program prices may not have declined steadily over time. In view of the institutional environment of the program and the predictions arising from models of dynamic competition with switching costs, it is possible to observe a nonmonotonic path for prices. Given the underlying complexities of the program design and the fact that consumer search was not exceptionally high in this market, the pattern the program prices followed is ultimately an empirical issue.

DATA

In this section we describe the drugs for which data were collected, the geographic areas covered, the timing of data collection, and the other prices obtained for control purposes.

Drugs

Prices were collected for 28 prescription drugs, which were chosen based on the following three criteria. First, all the drugs were in the top 100 drugs in claims filed by the elderly in 2001, and in the top 200 highest-selling drugs for the elderly in 2003. This selection of relatively popular drugs ensures that each drug had sufficiently large demand. The relatively high demand for these drugs implies that the price dynamics we are seeking are likely to have been apparent and economically important. Second, half of the drugs are

Table 1

Drugs Used in the Empirical Analysis

Drug name	Typical usage duration	Total sales rank (2003)	Rank in claims by elderly (2001)	Typical indications	Generic available?	Typical dosage	Among top 50 drugs for elderly (2001)?
Lipitor	Long term	1	5	Cholesterol	No	10 mg	Yes
Zocor	Long term	2	12	Cholesterol	No	20 mg	Yes
Norvasc	Long term	13	2	Cardiovascular	No	5 mg	Yes
Zoloft	Long term	7	27	Depression	No	50 mg	Yes
Lanoxin	Long term	NA	4	Cardiovascular	Yes	0.125 mg	Yes
Plavix	Long term	12	10	Cardiovascular	No	75 mg	Yes
Isosorbide mononitrate	Long term	NA	20	Cardiovascular	Yes	60 mg	Yes
Pravachol	Long term	18	38	Cholesterol	No	20 mg	Yes
Atenolol	Long term	175	45	Cardiovascular	Yes	25 mg	Yes
Metoprolol	Long term	NA	28	Cardiovascular	Yes	50 mg	Yes
Glucophage	Long term	99	9	Diabetes	Yes	500 mg	Yes
Detrol	Long term	86	32	Urinary	No	1 mg	Yes
Glucotrol XL	Long term	127	40	Diabetes	No	10 mg	Yes
Zestril	Long term	NA	33	Cardiovascular	Yes	10 mg	Yes
Amoxicillin	Short term	NA	> 100	Antibiotics	Yes	500 mg	No
Augmentin	Short term	177	> 100	Antibiotics	Yes	500 mg	No
Zithromax	Short term	NA	> 100	Antibiotics	No	500 mg	No
Minocycline	Short term	NA	> 100	Antibiotics	Yes	100 mg	No
Levaquin	Short term	25	> 100	Antibiotics	No	500 mg	No
Carisoprodol	Short term	NA	> 100	Pain	Yes	350 mg	No
Cephalexin	Short term	171	> 100	Antibiotics	Yes	250 mg	No
Ambien	Short term	31	> 100	Insomnia	No	10 mg	No
Cipro	Short term	48	> 100	Antibiotics	No	500 mg	No
Biaxin	Short term	138	> 100	Antibiotics	No	500 mg	No
Skelaxin	Short term	132	> 100	Pain	No	400 mg	No
Flexeril	Short term	NA	> 100	Pain	Y es	10 mg	No
Cefzil	Short term	152	> 100	Antibiotics	No	500 mg	No
Doxycycline hyclate	Short term	NA	> 100	Antibiotics	Yes	50 mg	No

Table 2

Variables Used in the Empirical Analysis

Variable	Description
LONG_TERM	Dummy variable, 1 if the drug is a maintenance drug, 0 if the drug is primarily for short-term use
GENERIC	Dummy variable, 1 if the drug has a generic equivalent or is itself generic, 0 if the drug is brand name
PRES_2003	The total number of prescriptions for a drug in 2003
PAT_EXPIRE	Dummy variable, 1 if the drug's patent had expired by 2004, 0 otherwise
PAT_EXCLUSIVE	Dummy variable, 1 if the drug has an exclusive patent for a specific condition, 0 otherwise
FDA_YEAR	The year a drug was approved by the FDA
WALGREENS	Dummy variable, 1 if the pharmacy is a Walgreens store, 0 otherwise
CVS	Dummy variable, 1 if the pharmacy is a CVS store, 0 otherwise
ECKERD	Dummy variable, 1 if the pharmacy is an Eckerd store, 0 otherwise
GEO	Dummy variable, 1 if the card offers national coverage, 0 otherwise
FEE	The fixed one-time enrollment fee to a given card in dollars
MFG	The number of manufacturers with which a card has a contract for discount prices
ASSIST	Dummy variable, 1 if the card offers enrollment assistance, 0 otherwise
MAIL	Dummy variable, 1 if the card has a mail-order option for drugs, 0 otherwise
FORMULARY	Dummy variable, 1 if the drug offers the entire formulary of Medicare-approved drugs, 0 otherwise
FRAC65+	Fraction of people ≥65 years or older in a zip code
MEDHINC	Median household income in a zip code
RENT	Median rent for renter-occupied housing units in a zipcode
FRACWHITE65+	Fraction of people ≥ 65 years in a zip code who are white
FRACFEM65+	Fraction of people ≥ 65 years in a zip code who are female
POP65+	Population in a zip code ≥ 65 years
POPWHITE65+	Population in a zip code ≥ 65 years and white
POPFEM65+	Population in a zip code ≥ 65 years and female

SOURCE: FDA, U.S. Food and Drug Administration.

short-term drugs, such as antibiotics and pain-killers, and the other half are long-term, maintenance drugs, such as those used for diabetes and cardiovascular diseases. The evolution of short-term drug prices is expected to differ from that of maintenance drugs, for which consumers are likely to search more intensely for a bargain. Finally, drug dosages were selected to reflect the most frequently prescribed dosages for the drugs, so that the demand is large relative to what it would be with unusually high or low dosages.[5] Each drug price pertains to a 30-day supply. The drugs and some of their attributes are presented in Table 1.

Geographic Areas

The price data from the MDDCP website were listed at the level of zip codes. Ninety zip codes were chosen by a random stratified sampling, designed to oversample zip codes with a

[5] Drug-specific information was obtained from *Mosby's Drug Consult* (2004, 2005), which features information on the typical usage and dosages of drugs.

greater proportion of the population composed of elderly residents, defined as individuals who are 65 years of age or older. To determine any demand side effects on prices, we needed to ensure a sufficient variation in market size and other demand shifters, such as income, for discount drugs. The population of residents 65 or older in a zip code is a proxy for the local market size for MDDCP cards. The proportion of elderly people in a zip code population varies in our sample from a low of 3 percent to a high of 92.6 percent; the average is 28 percent and the standard deviation is 25 percent. We also gathered zip code–level demographic data from the U.S. Census Bureau's 2000 Zip Code Statistics to analyze the price effect of demand shifters such as income and race composition (Table 2).

The program's price search engine listed prices for all pharmacies within a circle of a certain radius whose center coincides with the center of the selected zip code area. The search engine allowed for a choice of four different radii for any given zip code, and these radii varied by zip code. For densely populated urban areas, radii tended to be much smaller, whereas for less densely populated suburban and rural areas, the radii were larger, so that cardholders in these areas could obtain price information for a sufficient number of pharmacies. We collected price data for all pharmacies within the smallest and the second-smallest radii around a given zip code. This selection enabled us to assess the sensitivity of our results to the choice of radius.

Timing of Data Collection

The price data were updated weekly on the Medicare website between April 29, 2004, and December 31, 2005. As shown in Figure 1, the sample in this paper was collected for several weeks to cover important periods when the MDDCP was in effect.[6] Prices were first made available online on April 29, 2004, card enrollment began on May 3, 2004, and cards went into effect on June 1, 2004. Data collection was initiated on June 21, 2004, three weeks after subscribers were first allowed to use their cards under the program.[7]

The first wave of data was collected each week for a period of seven weeks during the summer of 2004. We refer to this period as the preswitching period. The second wave was collected during the last week of December 2004. This week is within the period between November 15 and December 31, which was the nationally coordinated switching period. Price observations from this period enable us to test whether card sponsors lowered their prices in an effort to induce switches. Finally, the third wave was collected after the end of the switching period to assess the behavior of prices when switching cards was not allowed. We label this period the postswitching period. During the postswitching period, the collection process took place over nine collection cycles, which included data from March 7, 2005, through August 15, 2005.

Each price observation pertains to a drug sold by a pharmacy at a given location under the discount offered by a given card at one point in time. The prices are posted prices, not necessarily transaction prices. Transactions may have taken place at only a subset of the posted prices, and some cards may have had little or no sales for some drugs. Therefore, posted prices do not necessarily coincide with the set of prices at which transactions take place. Lacking sales data, we are unable to make any statements on these issues. No card sponsor imposed explicit restrictions on the geographic or time variation in prices.[8] Geographic variation may have occurred for several reasons, including the changing demand and cost conditions or simply the changing composition of cards across locations.

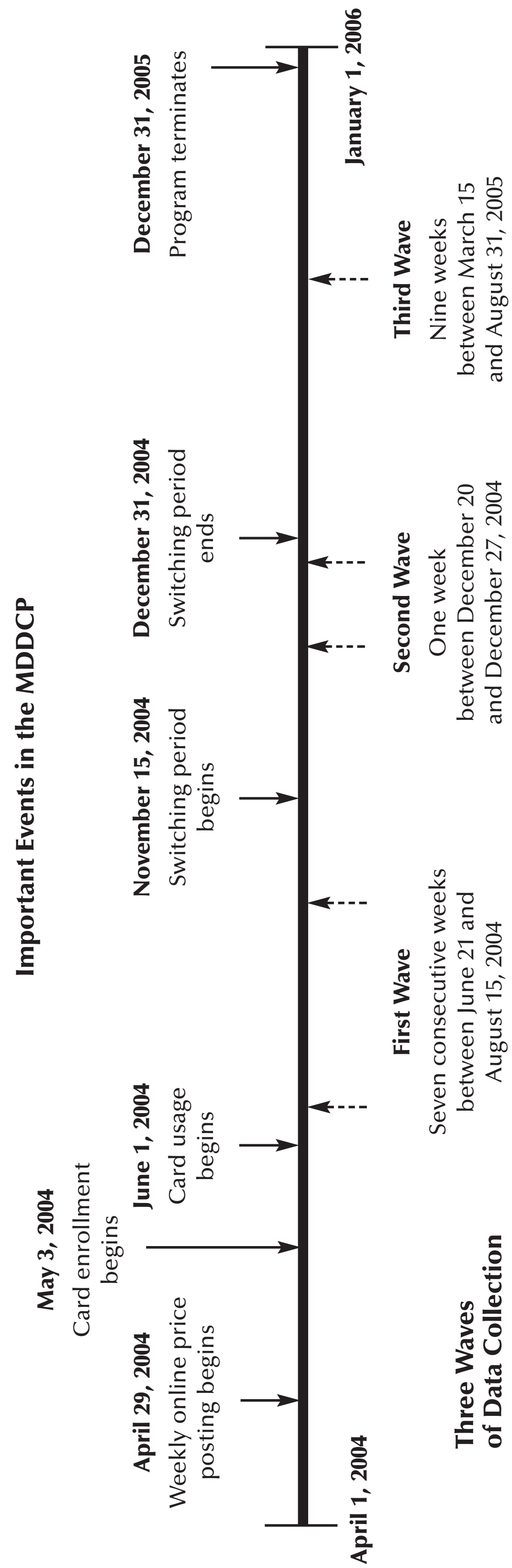
[6] The data collection process was automated using iOpus Internet Macros software that allowed periodic recording of the data from the Medicare website.

[7] The price data during the initial weeks of the program contained certain glitches, as noted by others (see Antos and Pinell, 2004). Some prices reported by pharmacies were found to be inaccurate and incorrect. However, these problems were fixed to a large extent within the first few weeks of the program. To ensure reliable data, we started collection in the fourth week after the cards went into effect.

[8] A brochure offered by a Walgreens store in Houston, Texas, specifically stated that prices were subject to change from store to store and over time.

Figure 1

Chronology of Important Events and Data Collection



Other Price Data

Part of our analysis aims to assess the magnitude of savings offered through card usage by controlling for changes in the general level of drug prices unrelated to the MDDCP. Ideal control data for this purpose would be comparable pharmacy-level, nonprogram retail price data collected at a weekly frequency to match the sample of MDDCP prices. Unfortunately, such detailed data are difficult to find. Instead, we collected nationwide wholesale prices for the drugs in our sample. The prices listed here are from *Mosby's Drug Consult* (2004, 2005), which provides prices for major drug wholesalers by dosage and duration. They are a representative sample of the wholesale prices typically used to reimburse patients for their prescriptions.[9] Unlike the card prices, however, these prices are not available by geographic units. Rather, a single nationwide price is reported by each supplier, usually a manufacturer. In addition, the price quotes are not available at a weekly frequency. Instead, they are representative of the price levels for the year the database was formed. Despite their shortcomings, these prices are the best readily available benchmarks and can be used to approximate savings. As we will show, the MDDCP prices exhibit little or no geographic dispersion. Thus, the nationwide prices in *Mosby's Drug Consult* can serve as a reasonable benchmark.

To attribute the evolution of prices to program specifics, the general trends exhibited by drug prices over the course of the program also need to be eliminated. For this purpose, we collected concurrent weekly prices posted by Internet drug retailers for the same drugs and dosages as in the program data. We used a major Internet prescription drug search engine, which quoted prices from several Internet drug retailers.[10]

Unlike program prices, online prices exhibit no geographic variation. Because they are subject to general nationwide trends in drug prices, they can serve as a good comparison group for the prices posted by card sponsors. The purpose of this comparison is twofold: First, it allows us to assess whether consumers would be able to obtain lower prices simply by purchasing at regular online prices available to general consumers, rather than going through the complicated process of choosing a card and hunting for lower prices. Second, and more importantly, online prices can be used to control for general changes in drug prices unrelated to the program. Drug prices can change over time because of changes in manufacturers' costs, availability of new substitute drugs, general inflation, or other factors. All such general trends are expected to apply in similar ways to MDDCP prices and online prices. Therefore, if different time patterns are observed for program prices versus other online prices, it is likely that program effects are an important cause. However, online prices may not reflect the exact set of nonprogram prices available to Medicare-eligible consumers. These consumers typically do not buy at regular online prices. Thus, online prices should not be viewed as an exact control group for Medicare-eligible consumers, but rather as a benchmark to control for general trends.

ANALYSIS

We begin with an analysis of the variation in price levels, followed by estimates of the extent of savings possible through the MDDCP. We then focus on price dynamics using the second-smallest radius for each zip code. The results were very similar when the smallest radius was used instead.

Analysis of Price Variation

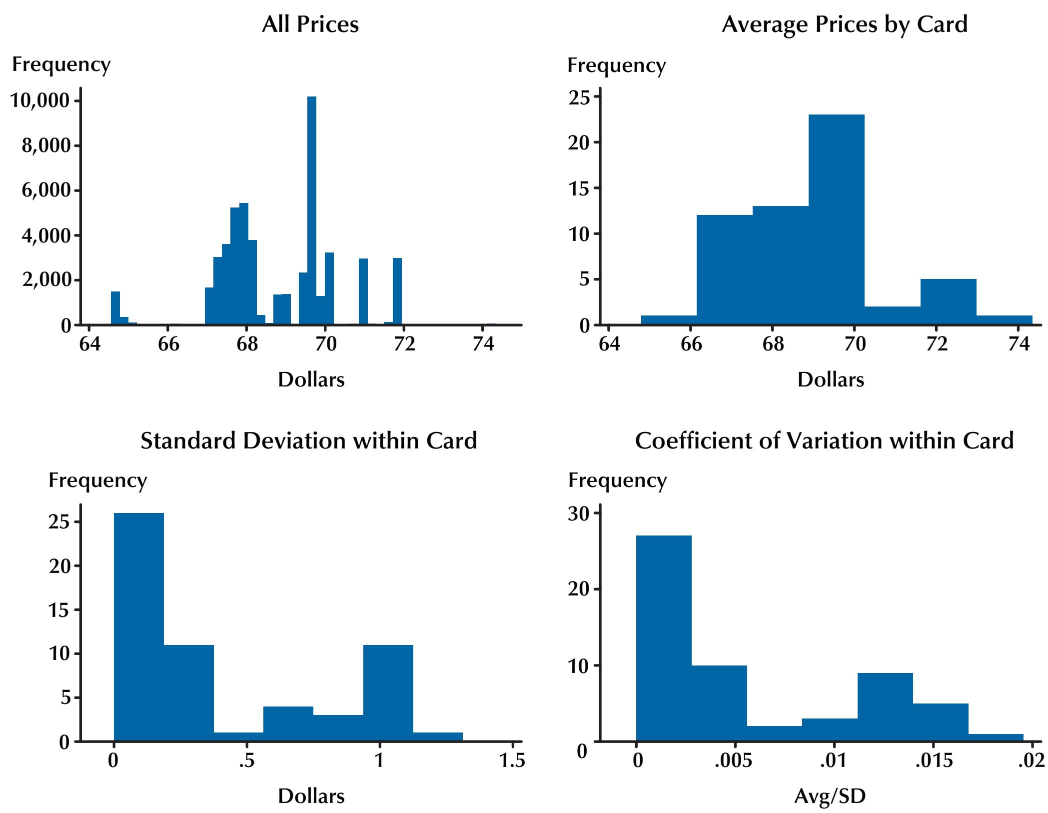
The starting point of our analysis is understanding whether significant price dispersion

[9] The nature of these prices is described in Mosby's reference book as follows: "Prices are AWP (average wholesale price), a benchmark price used for reimbursement. AWP represents what a retail pharmacist or a dispensing physician might pay for a product, without any special discounts. There are, however, many discounts already in place, so the AWP can often approximate the price that a consumer might pay. The prices listed here are not intended to serve as an up-to-date substitute for supplier price lists. The price listings give the reader a good idea of the range between the high and low prices."

[10] Once again, the data were collected by using iOpus Internet Macros from the website destinationrx.com. Our sample includes eight online retailers. Online stores of two major discount retailers (costco.com and walmart.com), online stores of two large drug retail chains (cvs.com and homemed.com), and the pharmacy branch of one major health care service provider (aarpharmacy.com).

Figure 2

Lipitor Price Dispersion: Week of June 28 to July 3, 2004



existed in the market for drug discount cards and, if so, what drove that dispersion. Figure 2 illustrates the dispersion of prices for one drug, Lipitor, for the week of June 28–July 3, 2004. The upper-left panel is the histogram of the entire set of Lipitor prices observed across cards, zip codes, and pharmacies. The upper-right panel is the distribution of the average price within a given card. The average price for a card is calculated using all price observations pertaining to the card. The average price varies between about \$65 and \$74. However, as shown in the two lower panels, the dispersion of price within a card is usually very small, amounting to an economically negligible variation across pharmacies within a card, even though such lack of variation was not explicitly guaranteed a priori by any card.

To determine whether the pattern in Figure 2 is typical of all drugs, we consider a general expression for the price p_{drczt} of drug d offered by card c at pharmacy r in zip code z at time t:

(1) $$p_{drczt} = \mu + f_d + f_r + f_c + f_z + f_t + e_{drczt},$$

where μ is a constant, f_i is a fixed effect for i $\in \{d,r,c,z,t\}$, and e_{drczt} is a zero mean error term that accounts for remaining unobserved factors. The contribution of each of the main factors to the overall variation in price can be analyzed by using analysis of variance (ANOVA) to understand the components of variation in prices. Because pharmacies are "nested" within zip codes, a nested ANOVA was performed to decompose the total variation in prices for each drug. Results of the ANOVA for the first week of data (June 21-27,

2004) showed that the variation in price of any drug across cards is the major component of the total variation in drug prices. On average, about 87 percent (standard deviation [SD] 15.3 percent) of the total variation in price is explained by the variation across cards, and there is little variation within cards. The variation across zip codes was only 0.5 percent (SD 2.1 percent) of the total variation on average, and the variation across pharmacies accounted, on average, for only 1.3 percent (SD 1.8 percent) of the total variation. The hypothesis that the average price of a drug is equal across cards is rejected strongly for all drugs. We repeated the ANOVA for other weeks and the findings supported the same conclusions.[11]

The finding of little variation in retail prices across zip codes raises the issue of how much geography affects pharmacies' pricing behavior. By the effect of geography, we mean the location-specific factors that may affect prices, such as income level of residents, population, and age composition in a location, which are particularly relevant as demand shifters. The ability to control for all other factors is important in investigating geographic variation in prices. The ideal experiment would look at the geographic variation in prices for a given drug and card combination, holding constant the pharmacy composition across zip codes. Such an experiment is impossible, however, because pharmacy composition changes across zip codes. Nevertheless, a close approximation to this ideal experiment is possible by looking at the prices charged by the stores of a given pharmacy chain across zip codes. The individual stores of a chain, such as Walgreens or CVS, tend to have very similar structures and practices, so a good approximation can be obtained by assuming that the store-level features are roughly constant across zip codes for a given chain. We calculated the coefficient of variation of prices across all stores of a pharmacy chain for each drug and card combination. In almost all cases, the coefficient of variation was either exactly zero or very close to zero. Thus, price variation across zip codes arose mainly because the composition of cards and pharmacies changed across zip codes.

Although the analysis of variance in prices clearly indicates that much of the cross-sectional variation was attributable to the variation across cards, it does not provide information about specific factors responsible for this variation. Identifying key demand and supply factors that affect prices is important for understanding why prices differed across drugs, cards, pharmacies, or zip codes.

Consider the following version of equation (1) that includes explanatory variables explicitly for a given a time period (week):

$$(2) \quad p_{drcz} = \mu + \beta_d X_d + \beta_r X_r + \beta_c X_c + \beta_z X_z + \varepsilon_{drcz},$$

where β_i is a $K_i \times 1$ vector of coefficients and X_i is a $K_i \times N$ matrix of observables, for $i = d,r,c,z$. Each X_i has the form

$$\left[x_1 x_2 \ldots x_{K_i}\right]',$$

where x_j is an $N \times 1$ vector that contains variables specific to cluster $j = 1,\ldots,K_i$ within group $i = d,r,c,z$.

The structure of the error term in equation (2) is assumed to be

$$(3) \qquad \varepsilon_{drcz} = \varepsilon_d + \varepsilon_r + \varepsilon_c + \varepsilon_z + e_{drcz},$$

where e_{drcz} is the error term in equation (1) and is assumed to be uncorrelated across observations. The error terms ε_i ($i \in \{r,d,c,z\}$) represent the remaining unobserved part of the fixed effect f_i in (1) after the observable X_i is added to the specification in equation (1) to obtain equation (2). Specification (3) implies that error terms are correlated within drugs, cards, pharmacies, and zip codes because of the presence of cluster-specific errors ε_i ($i \in \{r,d,c,z\}$). Because the component ε_i is fixed within cluster i, we can include dummy variables for drugs, cards, pharmacies, and zip codes to account for these unobserved components. The error term defined in equation (3) then reduces to e_{drcz} as in equation (1), which is assumed to be uncorrelated across observations within a cluster, and we can implement the regres-

[11] We also performed the ANOVA for mail-order prices for cards with a mail-order option. Not surprisingly, the entire variation in the case of mail-order prices (excluding shipping charges) was attributable to the cards.

Table 3

Static Price Regression

Independent variables	Dependent variable: Price I	Dependent variable: Price II
LONG_TERM	–51.10 (0.04)	–51.10 (0.04)
GENERIC	–11.58 (0.07)	–11.58 (0.07)
PRES_2003	–0.0000027 (0.0000001)	–0.0000027 (0.0000001)
PAT_EXPIRE	48.09 (0.06)	48.09 (0.06)
PAT_EXCLUSIVE	192.93 (0.09)	192.93 (0.09)
FDA_YEAR	1.89 (0.02)	1.89 (0.02)
WALMART	0.14 (0.05)	0.16 (0.06)
CVS	–0.94 (0.06)	–0.94 (0.06)
ECKERD	0.69 (0.03)	0.68 (0.04)
GEO	4.94 (0.61)	5.14 (0.61)
FEE	0.07 (0.01)	0.08 (0.01)
MFG	–0.44 (0.06)	–0.47 (0.06)
ASSIST	–4.15 (0.7)	–4.17 (0.7)
MAIL	2.04 (0.48)	2.13 (0.48)
FORMULARY	1.82 (0.15)	1.73 (0.17)
FRAC65+	–0.33 (0.09)	—
MEDHINC	–0.00037 (0.000013)	–0.00022 (0.000097)
RENT	0.0064 (0.00037)	0.0066 (0.00039)
FRACWHITE65+	–0.29 (0.03)	—
FRACFEM65+	–0.24 (0.02)	—
POP65+	—	–0.03 (0.002)
POPWHITE65+	—	0.029 (0.0017)
POPFEM65+	—	0.0024 (0.0003)
Card dummies	Y	Y
Drug dummies	Y	Y
Zip code dummies	Y	Y
Pharmacy dummies	Y	Y
N	1,230,215	1,230,215
R^2	0.98	0.98

NOTE: Robust standard errors are listed in parentheses.

sion in equation (2) without using any cluster effects.[12]

The results of the regression are shown in Table 3 for two specifications for the week of June 21-27, 2004. We used the same specification for other time periods and the results were robust. In evaluating the results, it should be noted that the drugs in our sample form only a subset of all drugs (28 of more than 800 drugs) covered by the MDDCP. Therefore, some characteristics that would apply in general to the drugs in the entire list of the MDDCP may not be fully represented in this relatively small sample.

The explanatory variables, including the dummies, account for 98 percent of the variation in prices. Given the large number of observations, almost all coefficients are precisely estimated. Long-term maintenance drugs in our sample were, on average, cheaper than the short-term drugs, based on the prices for 30-day supplies.[13] Generic drugs and brand-name drugs for which generic alternatives are available were cheaper compared with drugs that do not have generic alternatives. The prices were also lower for drugs that are prescribed more frequently. In addition, newer drugs had higher prices, as indicated by the positive coefficient on the year of approval by the FDA.

The coefficients on selected pharmacy chains suggest that Wal-Mart had slightly higher prices, by about 14 cents, than those of the omitted category of all remaining pharmacies, while CVS prices were lower by about a dollar than those of the omitted category. Eckerd, which merged with CVS in the spring of 2004 shortly before the MDDCP took effect, had prices that were higher by about 70 cents than those of the omitted category.

Cards with national coverage and with a mail-order service tended to have higher prices than cards that did not have national coverage and mail-order service. Cards with higher subscription fees and with a broader formulary also tended to have higher prices. Cards that had arrangements to provide discounts with a larger number of drug manufacturers and those that provided enrollment assistance had lower prices than the cards that did not offer these benefits. Certain quality dimensions, such as formulary breadth, extensive geographic coverage, and cost-reducing features such as association with a larger number of manufacturers, apparently were important for the differences in price across card sponsors.

Demographic characteristics of zip code areas also influenced prices to some extent. Zip codes with a higher proportion of elderly people in the population had lower prices. Zip codes with a higher median household income also had lower prices, whereas zip codes with higher housing costs were associated with higher drug prices; but these effects are relatively small.

Estimates of Savings

The finding of differences in prices across cards noted in the previous section raises the following question: Were the differences large enough to reward searching for lower prices across cards? Several small-scale studies tried to assess the extent of the discounts in the early phases of the program with only a handful of drugs and a few zip codes.[14] Such investigations generally found some savings accruing to cardholders, but the small scale of these investigations prevented any general conclusions. In the following text, we ignore the card enrollment fee, which in most cases was zero and could not exceed $30, and look only at the savings a cardholder could obtain from using the card to purchase drugs at card prices versus purchasing at regular retail or online prices.

[12] Because of the large number of dummy variables in this regression (>1,000 pharmacy dummies), we use the "de-meaned" regression approach (Greene, 1993, pp. 468-69). By de-meaning the observations by pharmacy, we eliminate the pharmacy dummies and still obtain the usual ordinary least square (OLS) estimates of the coefficients of interest.

[13] The price difference should not be taken as evidence that the cost of therapy is lower for long-term drugs in general, because long-term prescriptions typically are renewed for several months and some short-term prescriptions are prescribed for periods shorter than a month (antibiotics such as Zithromax, which are used for intense treatment for a week in certain cases). If the drug is used only seven days, the cost of therapy will be low.

[14] See, for example, Antos and Ximena (2004). Their approach first identifies a few health conditions that are common among the elderly and then calculates the total price of a bundle of drugs typically prescribed to remedy these conditions.

Let $\bar{p}_{dct}$ be the average price of drug d for card c in week t, where the average is taken across all pharmacies selling drug d offered by card c. Define $\bar{p}_{dt}$ and p_{dt}^{min} as the average and the minimum of $\bar{p}_{dct}$ across cards in week t. Similarly, define the average and minimum regular prices obtained from Mosby's (2004) database as $\bar{p}_{d}^{Mosby}$ and $p_{d}^{Mosby,min}$, where the average and the minimum are calculated across the wholesalers listed in the Mosby's database for a given drug. In addition to the prices in Mosby's database, a separate, independent source is the set of prices we collected from online pharmacies as described earlier. Define the average and minimum prices for online retailers in week t as $\bar{p}_{dt}^{online}$ and $p_{dt}^{online,min}$.

We now define several alternative measures of potential savings. The first measure is the savings a *naive* (or nonsearching, or uninformed) consumer could obtain. A naive consumer is defined as one who purchases randomly with equal probabilities across cards. For a single purchase of the drug at a given point in time, if this consumer uses a card instead of buying at the regular wholesale price (i.e., outside the discount program), the savings are the percentage difference between the average regular price and the average card price. We report the average of these savings across all weeks in the data in percentage form as follows:

$$(4) \qquad S_d^{naive} = \frac{100}{T}\sum_{t=1}^{T}\left(\frac{\bar{p}_d^{Mosby} - \bar{p}_{dt}}{\bar{p}_d^{Mosby}}\right).$$

The second measure is the savings that accrued to an Internet *searcher*, who uses the program website to search for the lowest-price card for a given drug, but otherwise would purchase randomly in the regular market (outside the program) because of higher search costs (as opposed to searching for discounted prices online). The savings of such a consumer are defined as the percentage difference between the average price in the regular market and the minimum price in the discount card market averaged across weeks, and are obtained simply by replacing $\bar{p}_{dt}$ in equation (4) by p_{dt}^{min}.

The third measure we consider is the savings an *expert* consumer could obtain. An expert consumer is defined as one who is fully informed of prices in both markets and thus is always able to purchase at the minimum price. The average savings across weeks for such a consumer are formally defined as the percentage difference between the minimum price in the regular market and the minimum price in the discount card market averaged across weeks, and are obtained by replacing $\bar{p}_d^{Mosby}$ in equation (4) by $p_d^{Mosby,min}$ and $\bar{p}_{dt}$ by p_{dt}^{min}.

Following Baye, Morgan, and Scholten (2003), we also define the "value of information" in the drug discount card market, which is the saving of a consumer informed of all card prices with respect to that of a naive consumer,

$$(5) \qquad V_d^{card} = \frac{100}{T}\sum_{t=1}^{T}\left(\frac{\bar{p}_{dt} - p_d^{min}}{\bar{p}_{dt}}\right).$$

We also report the value of information for online prices and prices from Mosby's database.

The defined savings measures and the values of information are reported by drug in Table 4. A naive consumer could obtain an average savings of 11.2 percent. The average savings were even higher for a searcher—about 25 percent. An expert consumer, on the other hand, had little to gain from purchasing in the discount card market: An average savings of only 2.3 percent accrued to such a consumer. Because most drug card users were likely non-experts in searching, the estimate of savings to naive consumers, or at best to searchers, is likely to be the most reasonable estimate.

A somewhat different picture emerges when we consider the savings with respect to online prices. A searcher could obtain an average savings of 8.7 percent by purchasing at the minimum card price instead of purchasing randomly from one of the online pharmacies. However, the benefit for a naive consumer was negative (but statistically insignificant), and an expert consumer could obtain positive (again statistically insignificant) savings. Thus, compared with online prices, card prices did not appear to provide substantial savings. The average value of information also was the highest for regular prices, indicating the biggest rewards to searching, with an average savings for an informed consumer that amounted to around 20 percent of the average price. These savings were

Table 4

Estimates of Savings from Drug Discount Cards

	Savings (percent)								
	Regular versus card prices			Online versus card prices (with shipping)*			Value of information		
Drug	Naive†	Searcher‡	Expert§	Naive*	Searcher‖	Expert	Online	Card	Regular
Ambien	5.2	17.5	14.2	12.0 (10.8)	13.6 (12.6)	0.7 (–0.4)	1.8 (2.0)	13.0	3.8
Amoxicillin	50.1	63.6	10.8	–42.0 (–66.3)	–16.9 (–27.0)	–54.4 (–70.5)	18.2 (22.4)	24.1	59.3
Atenolol	54.4	75.6	67.4	22.8 (9.7)	45.7 (45.7)	–1.7 (–1.7)	29.0 (39.0)	46.2	25.5
Augmentin	–1.6	8.3	–29.3	3.2 (2.6)	5.0 (3.8)	–5.1 (–6.5)	1.8 (1.2)	9.6	29.0
Biaxin	5.7	9.0	–22.0	–4.0 (–5.1)	3.4 (1.9)	–0.08 (–1.6)	7.2 (6.7)	3.5	25.4
Carisoprodol	38.7	89.4	43.5	48.6 (43.7)	57.3 (57.3)	–147.6 (–147.6)	16.8 (24.1)	82.6	81.2
Cefzil	14.3	16.5	–12.7	–19.1 (–19.4)	–17.3 (–17.8)	–20.2 (–20.7)	1.9 (1.6)	2.5	25.9
Cephalexin	NA	NA	NA	69.2 (66.2)	77.6 (73.2)	47.1 (36.7)	27.5 (20.8)	57.6	NA
Cipro	5.0	9.8	–12.2	5.2 (4.4)	8.8 (8.0)	4.0 (3.1)	3.8 (3.8)	5.0	19.6
Detrol	5.9	11.1	1.9	–1.39 (–3.2)	3.4 (–0.4)	–2.3 (–6.4)	4.7 (2.7)	5.6	9.4
Doxycycline hyclate	63.4	79.2	79.3	NA	NA	NA	NA	43.1	0.0
Flexeril	6.8	21.9	3.7	–4.2 (–7.8)	5.9 (2.6)	–31.7 (–37.3)	9.8 (9.5)	16.3	19.0
Glucophage	–25.1	–11.0	–137.6	NA	NA	NA	7.6 (11.5)	11.3	39.5
Glucotrol XL	–12.2	–3.6	–71.4	–31.2 (–39.9)	–29.8 (–37.3)	–41.6 (–49.8)	1.3 (2.4)	7.9	53.3
Isosorbide mononitrate	59.6	81.8	74.8	–19.2 (–27.2)	–19.2 (–27.2)	–222.3 (–244.2)	0.0 (0.0)	54.4	27.8
Lanoxin	–2.8	11.5	–38.9	4.0 (–9.4)	16.1 (16.1)	2.4 (2.4)	12.6 (23.2)	13.9	36.3
Levaquin	10.1	14.2	14.2	–1.4 (–1.8)	3.9 (3.2)	–0.8 (–1.5)	12.6 (23.2)	4.6	0.0
Lipitor	5.7	11.3	6.5	4.2 (2.1)	8.7 (7.1)	2.9 (1.2)	4.6 (5.0)	6.0	5.1
Metoprolol	NA	NA	NA	–23.6 (–48.8)	24.8 (24.8)	–23.9 (–23.9)	39.1 (49.5)	39.3	NA
Minocycline	–46.3	24.7	24.7	47.9 (46.6)	61.6 (61.6)	25.4 (25.4)	25.4 (27.2)	48.5	0.0
Norvasc	6.8	12.0	2.0	1.2 (–2.0)	6.5 (5.4)	1.1 (–0.09)	25.4 (27.2)	5.5	10.2
Plavix	–2.0	3.1	–17.0	2.1 (0.8)	5.5 (4.9)	0.5 (–0.02)	3.4 (4.2)	5.0	17.2
Pravachol	–9.5	–1.9	–1.9	–4.8 (–6.5)	1.2 (0.4)	–6.1 (–7.0)	5.6 (6.5)	6.9	0.0
Skelaxin	–16.4	–8.6	–30.9	–9.1 (–10.8)	–1.5 (–4.7)	–8.1 (–11.7)	8.3 (6.3)	6.6	17.1
Zestril	28.5	36.1	31.7	–31.6 (–35.8)	–25.5 (–29.2)	–44.4 (–48.7)	4.1 (4.3)	11.3	6.5
Zithromax	14.2	20.3	20.3	–2.6 (–2.9)	–2.6 (–2.9)	–10.4 (–10.8)	4.1 (4.3)	7.0	0.0
Zocor	24.4	42.9	34.1	–20.2 (–21.5)	–15.9 (–15.9)	–53.4 (–53.4)	0.0 (0.0)	24.4	13.4
Zoloft	8.0	14.0	5.3	3.0 (1.2)	6.9 (5.1)	0.5 (–1.6)	4.0 (3.9)	6.5	9.2
Average	**11.2#**	**24.9#**	**2.3**	**0.4 (–4.6#)**	**8.7# (–19.0#)**	**–22.6# (–26.0#)**	**9.1# (10.6*)**	**20.3#**	**20.5#**
SD	25.9	29.1	44.9	25.1 (28.1)	26.3 (20.6)	53.8 (57.0)	10.0 (12.4)	21.1	20.3
Median	**6.3**	**14.1**	**4.5**	**–1.4 (–3.0)**	**5.2 (–10.3)**	**–3.7 (–6.4)**	**5.0 (4.9)**	**10.5**	**17.1**
Interquartile range	23.7	24.1	39.5	20.8-19.7	14.7-23.1	32.4-37.2	7.7-13.6	22.2	22.7

NOTE: *The figures within parentheses include shipping fees. The figures outside the parentheses are based on the online base prices. †"Naive" is defined as a consumer who is uninformed and purchases randomly in both markets. ‡"Searcher" is defined as a consumer who is informed of the minimum card price but otherwise purchases randomly in the regular market. §"Expert" is defined as a consumer who is fully informed in both markets. ‖"Searcher" is defined as a consumer who is informed of the minimum online price but otherwise purchases randomly in the discount card market. #Indicates difference from zero at 5 percent or lower levels.

followed closely by card prices. The value of information in the online market was the lowest.

Dynamics of Prices

We now turn to the evolution of prices. Price changes using two balanced panels of pharmacies from the preswitching period and the postswitching period are examined in the first subsection below. Next, we investigate the behavior of prices around the switching period. The evolution of online prices is examined for comparison with program prices, followed by consideration of the evolution of price dispersion within the program.

Results from the Balanced Panels

Using a slight modification of equation (1), a price observation can be written as

$$(6)\quad p_{drczt} = \mu + f_t + f_{ct} + f_{dt} + f_d + f_c + f_r + f_z + \eta_{drczt},$$

where we introduced the interaction terms, f_{ct}, a card- and time-specific effect, and f_{dt}, a drug- and time-specific effect. The term f_{ct} captures potentially different behavior of cards over time. Different cards may have had different pricing policies that may have depended on time as competition among card sponsors changed. In addition, the time and drug interaction effect, f_{dt}, captures the possibility of different drugs experiencing different price changes over time (e.g., cards may have competed more intensely in certain popular drug categories). The fixed effect, f_t, can be interpreted as the general time effect on prices, which is a combination of the program's effect on price and general fluctuations in drug prices outside the program.

The specification in equation (6) can be estimated using our unbalanced panel of observations. This approach has two drawbacks. First, a very large number of effects (both pure and interaction effects) must be estimated. Second, and more importantly, the included effects are not guaranteed to exhaust the set of relevant effects, which may lead to omitted variable bias, and the time-invariant fixed effects can potentially be correlated with the error term. One approach to alleviate these concerns is to use time differencing, which eliminates the time-invariant fixed effects. By taking the difference of the prices for two consecutive time periods, t and t', we obtain

$$(7)\quad \Delta p_{drczt} = d_{ct} + d_{dt} + d_t + \varepsilon_{drczt},$$

where d_{ct}, d_{dt}, and d_t are the obvious time differences for the corresponding fixed effects and $\varepsilon_{drczt} = (\varepsilon_{drczt} - \varepsilon_{drczt})$. Because differencing works only if we have the same pharmacies across the two time periods, we restrict attention to a balanced panel.

Now, consider the following ordinary least squares (OLS) regression based on equation (6):

$$(8)\quad \Delta p_{drczt} = \beta_{ct} D_{ct} + \beta_{dt} D_{dt} + \beta_t D_t + \varepsilon_{drczt},$$

where D_{ct}, D_{dt}, and D_t are dummies for the differenced effects d_{ct}, d_{dt}, and d_t. The error term ε_{drczt} has serial correlation, which we take into account in estimating the standard errors.

One problem with this approach is that the balanced panel has a low cross-sectional dimension if we restrict attention only to observations common across all weeks of data in the sample period. Because of errors in accessing the MDDCP website that occurred randomly during the data collection, there was some attrition in our sample and the balanced panel that can be constructed across all weeks of observation is limited in size. However, because this attrition was entirely random, a systematic bias is not a concern. Consequently, we implement regression (8) separately for the seven weeks in the preswitching period and then for the nine weeks in the postswitching period. This approach provides a large, but different, number of cross-sectional observations for both periods. We handle the data for the switching period separately as discussed below.

We first consider the evolution of prices using a panel from weeks 4 to 10 of the program, the preswitching period. The results of the difference regression for this period are shown on the left side of Table 5. The estimates of β_t are all negative and statistically significant, except for week 5 of the program. Most of the drop in prices in this period took place between the fifth and eighth weeks, resulting in a decline in general level of prices of about \$4.77. By the end of the 10th week, the prices were lower by about \$4.63. However, this

Table 5

Estimated Coefficients of Time Dummies from the Difference Regressions

Dependent variable: first difference in price

Independent variables: dummy for the week of	Estimates for preswitching period	Independent variables: dummy for the week of	Estimates for postswitching period
6/28/2004	0.21 (0.27)	4/4/2005	2.85 (0.11)
7/5/2004	–2.25 (0.38)	5/16/2005	4.53 (0.16)
7/11/2004	–3.73 (0.47)	6/6/2005	6.54 (0.20)
7/18/2004	–4.77 (0.54)	6/20/2005	7.74 (0.23)
7/25/2004	–4.64 (0.61)	7/11/2005	7.74 (0.26)
8/2/2004	–4.63 (0.66)	7/18/2005	7.7 (0.28)
		8/1/2005	7.8 (0.31)
		8/15/2005	7.83 (0.33)
N	92,700	N	18,280
R^2	0.53	R^2	0.51

NOTE: Robust standard errors are shown in parentheses. Omitted time dummy is the first week for each regression: 6/21/2004 for the preswitching period and 3/7/2005 for the postswitching period.

reduction represents a small portion (5.5 percent) of the average ($81.90) of all price observations during the fourth week of the program when data collection began.

We repeated the analysis for the postswitching period using a balanced panel. The evolution of the prices in the sample of weeks from the postswitching period shows a different pattern compared with the preswitching period, as seen on the right side of Table 5. In fact, the estimated β_t coefficients are all positive and statistically significant. Between the starting and ending weeks of the sample in the postswitching period, prices rose by about $8, controlling for drug and card effects. Much of this increase took place between the end of the switching period and the end of June 2005. Thereafter, prices stabilized somewhat. Between the end of the switching period and the end of June, prices rose at a pace of about $2 a month. The total rise in prices represents about 9.5 percent of the average drug price in the fourth week of the program.

Figure 3 displays the discrepancy in the average evolution of prices for different cards and drugs. Specifically, the upper two histograms display the frequency distributions of the time average of the card-time effects plus the pure-time effects,

$$\bar{\hat{\beta}}_c = \frac{1}{T}\sum_{t=1}^{T}\left(\hat{\beta}_t + \hat{\beta}_{ct}\right), \tag{9}$$

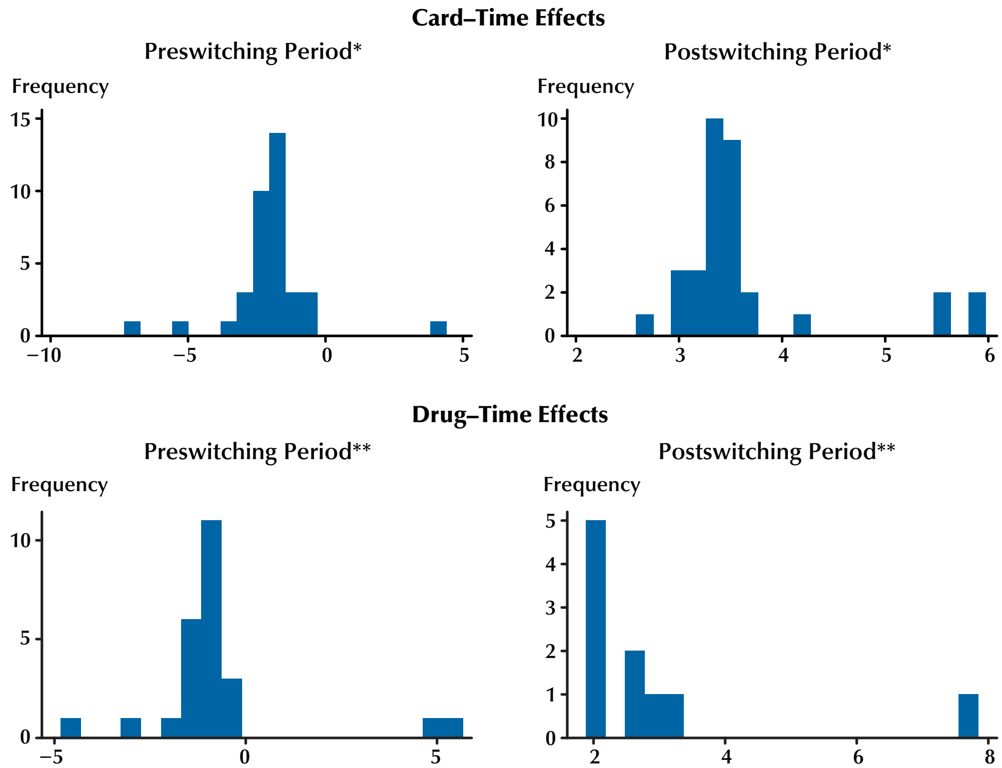
for the preswitching and postswitching periods, on the left and the right panels, respectively. Analogously, the bottom two panels contain the frequency distributions of the time average, of the drug-time plus the pure-time effects,

$$\bar{\hat{\beta}}_d = \frac{1}{T}\sum_{t=1}^{T}\left(\hat{\beta}_t + \hat{\beta}_{dt}\right), \tag{10}$$

for the preswitching period on the left and the postswitching period on the right. As is evident from the histograms on the left-hand side of the upper and lower panels, most cards and drugs had lower prices during the preswitching period. However, the right tails of these histograms show a few outlier cards and drugs that exhibited an average upward trend in prices even during this period. In contrast, for the postswitching period, all cards and drugs exhibited an average upward

Figure 3

Frequency Distribution of Average Card–Time and Drug–Time Effects



NOTE: *As defined in equation (9). **As defined in equation (10).

trend in price (seen in the histograms on the right-hand side of the upper and lower panels). Overall, these histograms suggest that prices of cards and drugs on average moved in the same direction within the preswitching and postswitching periods with few exceptions.

We repeated the estimation in equation (8) by adding a long-term drug dummy interaction with a time dummy to explore whether long-term drugs exhibited any different behavior compared with short-term drugs. We found that during the preswitching period, the prices for long-term drugs actually fell less, and during the postswitching period they rose less compared with short-term drugs. This pattern does not support the hypothesis that consumers searched more vigorously for bargains on these drugs. If this were the case, we would have expected to see a steeper decline for these prices compared with the prices of short-term drugs.[15]

The Switching Period

For the nationally coordinated card-switching period between November 15 and December 31, 2004, we were able to collect only one week of price data because of technical problems in accessing the website during much of that period. As a result, we were able to collect data for only 15 drugs, and the generally smaller number of observations for that period precluded us from including the switching period in the balanced

[15] One possible explanation is that consumers with an existing prescription for a given long-term drug who have purchased from their preferred pharmacy for a long time may not have found it worthwhile to search vigorously for a card and a possibly different pharmacy—which illustrates another form of switching costs.

Table 6

Analysis of Price Changes Around the Switching Period

Drug	(Switching period price) – (Preswitching period price) Average difference ($)	Paired t statistic	p Value	(Postswitching period price) – (Switching period price) Average difference ($)	Paired t statistic	p Value
Ambien	–1.99	**–6.91**	0.00	0.31	**5.42**	0.00
Amoxicillin	–2.28	**–26.56**	0.00	1.33	**11.38**	0.00
Atenolol	–0.77	**–9.62**	0.00	0.18	**0.67**	0.49
Augmentin	–3.07	**–2.46**	0.01	0.72	**2.88**	0.00
Biaxin	–2.65	**–10.55**	0.00	2.80	**22.66**	0.00
Carisoprodol	–0.63	**–1.29**	0.04	1.50	**3.29**	0.00
Cefzil	–2.93	**–17.94**	0.00	–0.10	**–0.89**	0.37
Cipro	–4.81	**–5.00**	0.00	3.69	**3.35**	0.00
Detrol	–2.09	**–4.18**	0.00	3.27	**12.47**	0.00
Doxycycline hyclate	0.58	**8.73**	0.00	0.21	**4.25**	0.00
Flexeril	–0.70	**–3.18**	0.00	2.61	**16.08**	0.00
Glucotrol XL	0.34	**7.92**	0.00	1.03	**14.00**	0.00
Isosorbide mononitrate	–3.36	**–14.22**	0.00	2.30	**1.31**	0.18
Lanoxin	1.32	**22.87**	0.00	0.05	**0.60**	0.54
Levaquin	–3.80	**–10.78**	0.00	2.48	**12.42**	0.00
Average	–1.79			1.49		
Standard error	0.45			0.33		

NOTE: "Switching period price" is the price during the one week of data available from the switching period. "Preswitching period price" is the price during the last week (week of 8/2/2004) of price observations in our preswitching period sample. "Postswitching period price" is the price during the first week (week of 4/4/2005) of price observations in our postswitching period sample. Bold t statistics indicate significance at 5 percent or lower levels.

panel analysis of the previous section. Instead, we compared the average price level for each drug using two paired t tests. For each drug, we perform two paired t tests across common cards and pharmacies: one for the difference between the week from the switching period and the last week of the preswitching period, and the other for the difference between the first week of the postswitching period and the week from the switching period. The paired t test approach eliminates the fixed effects that are common across the two periods and isolates the time effects, just like the balanced panel used previously.

As shown in Table 6, both tests indicated a statistically significant decline in prices for most drugs (12 of 15) between the last week of the preswitching period and the week of the switching period, and a subsequent statistically significant rise for most drugs (11 of 15) between the week of the switching period and the first week of the postswitching period. The magnitude of price drops and raises varied across drugs. Overall, prices declined on average by about $1.80 between the week of August 2, 2004, and the week of December 20, 2004, and rose on average by about $1.50 between the week of December 20, 2004, and March 7, 2005.

Given the nature of the timing of data collection, we cannot say precisely whether the decline in prices between the week of August 2, 2004, and the week of December 20, 2004, was confined to the switching period only. Because card enrollment continued during this period, card sponsors could have continued to reduce their prices to some extent to attract additional consumers, as they did in the initial phases of the program. Some card sponsors, in anticipation of the switching period, may have also lowered prices in an effort to deter consumers from switching. Thus, some of the observed price decline in this period could have occurred even before the switching period. We are more comfortable attributing the rise in prices after the switching period to the existence of switching costs, because during that period card enrollment diffused to a large extent and enrolled consumers were committed until the end of the program.

In summary, the evidence from the balanced panel estimation and the paired *t* tests points to initially declining but later rising prices, even though the magnitudes of change in price levels were not exceptionally large compared with the average price level across drugs. The pattern exhibited by prices lends more support to a model in which prices move in a nonmonotonic path, falling when consumers could switch cards and rising when they could not switch cards.

Evolution of Nonprogram Online Prices

We now consider the evolution of online drug prices as a benchmark for the evolution of program prices. If the time effects found in the evolution of program prices are specific to the program rather than being driven entirely by general trends in drug prices, the same time effects should not emerge for online prices unrelated to the program. To explore this possibility we consider a regression of the form

$$(11) \qquad p_{dit} = \alpha + \beta_t D_t + \beta_i D_i + \beta_d D_d + \varepsilon_{dit}$$

where D_t is a time dummy, D_i is a dummy for online retailer *i*, and D_d is a dummy for drug *d*. The focus is once again on the estimates of the coefficients of time dummies.

Few problems were encountered in data collection of online prices over time, so our sample includes a larger number of weeks and the price changes can be observed with a higher frequency over a longer period, sometimes even more frequently than once a week. Table 7 presents the results of the estimation in equation (11). The time dummies have almost uniformly positive and significant coefficients, and the coefficients are almost monotonically increasing over time. By the last week of data, prices were higher by about \$3.39, controlling for vendor and drug fixed effects.

We repeated the estimation in equation (11) using the total price (base price plus shipping fee) as the dependent variable and the results were very similar. The total price increased over time by about \$3.53 and the estimated coefficients were uniformly positive and statistically significant in almost all cases.

Finally, we also used a balanced panel approach as in equation (8) to estimate the time effects for online prices. The size of this panel was much smaller than that of the unbalanced panel used in equation (11), because we did not have prices for all sellers and for all drugs every week. The average growth rate of price between the first and the last periods of observation was 3.31 percent (SD 0.11 percent). Only four drugs exhibited a decline in price. Overall, the results from the balanced panel were similar qualitatively to the estimates of time dummy coefficients in Table 7.

The observed pattern for online drug prices thus indicates that the evolution of program prices was indeed different from the evolution of prices outside the program. Online prices tended to rise over time, in contrast to the program prices, which first declined and later increased. Because online prices are subject to general trends in drug prices, but not to the effects of the program, the patterns suggest that the evolution of program prices is at least in part driven by program effects, rather than entirely by general trends. First, online prices rose during the preswitching period when the program prices exhibited a clear decline. The decline in program prices is consistent with the predictions of dynamic price competition models, suggesting an escalated competition in the early stages of a market when sellers lower their prices to lure consumers. Second, the overall rise in online prices fell short of the rise in program prices during the postswitching period.

Table 7

Estimated Time Dummies for Online Price Regression

Independent variables	Dependent variable		Independent variables	Dependent variable	
Dummy for the date:	Base price	Total price	Dummy for the date:	Base price	Total price
Preswitching period			**Postswitching period**		
6/28/2004	0.00 (0.58)	0.00 (0.58)	1/13/2005	1.95 (0.63)	1.99 (0.63)
7/8/2004	0.00 (0.58)	0.00 (0.58)	5/6/2005	2.62 (0.62)	2.66 (0.62)
7/15/2004	0.00 (0.58)	0.00 (0.58)	5/27/2005	2.81 (0.62)	2.85 (0.62)
7/26/2004	1.73 (0.60)	1.74 (0.60)	6/10/2005	2.81 (0.62)	2.85 (0.62)
8/3/2004	1.73 (0.60)	1.74 (0.60)	6/20/2005	2.81 (0.62)	2.85 (0.62)
8/10/2004	1.73 (0.60)	1.74 (0.60)	7/11/2005	3.25 (0.62)	3.31 (0.62)
8/17/2004	1.73 (0.60)	1.74 (0.60)	7/29/2005	3.25 (0.62)	3.31 (0.62)
8/24/2004	1.73 (0.60)	1.74 (0.60)	8/1/2005	3.30 (0.62)	3.36 (0.62)
9/1/2004	1.81 (0.61)	1.87 (0.61)	8/18/2005	3.33 (0.66)	3.38 (0.66)
9/13/2004	1.82 (0.62)	1.87 (0.62)	9/16/2005	3.39 (0.66)	3.53 (0.66)
9/15/2004	1.83 (0.61)	1.85 (0.61)	9/29/2005	3.39 (0.66)	3.53 (0.66)
9/21/2004	1.86 (0.60)	1.91 (0.61)	10/4/2005	3.39 (0.66)	3.53 (0.66)
9/24/2004	1.86 (0.60)	1.91 (0.61)	10/16/2005	3.39 (0.66)	3.53 (0.66)
9/28/2004	1.86 (0.60)	1.91 (0.61)	10/17/2005	3.39 (0.66)	3.53 (0.66)
10/5/2004	1.81 (0.63)	1.85 (0.63)	10/20/2005	3.39 (0.66)	3.53 (0.66)
10/15/2004	1.88 (0.63)	1.89 (0.64)			
10/20/2004	1.55 (0.64)	1.56 (0.64)			
Switching period					
12/10/2004	1.93 (0.62)	1.97 (0.62)			
12/29/2004	1.89 (0.62)	1.93 (0.62)			
N	2,955				
R^2	0.98				

NOTE: Robust standard errors are shown in parentheses. Total price includes shipping fee for standard delivery for each vendor. Dates refer to the day the price data were collected. Preswitching period: before 12/2004; switching period: 12/2004; postswitching period: after 12/2004.

Indeed, the program prices actually increased about $4 more than online prices by the end of this period. Therefore, the upward trend in program prices after the switching period cannot be explained simply by a general rise in drug prices caused by nonprogram effects.

In addition to the evolution of levels of prices, we also investigated the evolution of price dispersion. To measure price dispersion at any point in time, we first calculated the average of a drug's price within a card. Next, we computed the dispersion of that average around its mean across cards. We then used the balanced panel of observations to test the hypothesis that the price dispersion remained the same over time versus the alternative that dispersion changed. We found no overwhelming evidence that the dispersion of average price across cards changed substantially during the preswitching or the postswitching periods. The dispersion of prices was persistent over the course of the program.

CONCLUSION

We used a large panel of drug prices to assess the effects of government-sponsored release of price information over the Internet under the MDDCP. The designers of the program began with the premise that access to price information by consumers would lower prices over time. In contrast, the card prices and their dispersion did not steadily decline over time, as some models of improved access to price information suggest. Instead, prices declined during the initial phases of the program but then increased later when consumers were unable to switch cards. The evolution of program prices exhibited significant deviation from the general evolution of prices outside the program.

The nonmonotonic evolution of program prices can be reconciled with the predictions of certain models of dynamic price competition with consumer switching costs, such as that of Klemperer (1987). The very design of the program left consumers vulnerable to price changes by card sponsors. Card sponsors appeared to have reduced their prices initially, possibly in an effort to lure customers to subscribe, but then raised their prices in the later stages of the program to take advantage of consumers when they were locked in to their choices. However, we are unable to provide any direct evidence on the actual subscription patterns by card or whether consumers switched cards at all.

The extent to which these patterns will carry over to Medicare's Part D prescription drug assistance program currently in effect remains to be seen. Although Part D has a much more complicated structure, some drivers of price dynamics under MDDCP also apply to Part D. For instance, consumers can switch plans only from November 15 through December 31 of every year, except in special cases. There are also certain differences between the two programs. Consumer non-enrollment in Part D carries a financial penalty that becomes gradually more severe, unlike in the case of the MDDCP, where enrollment was voluntary. Also, the prescription drug benefit providers engage in a multiperiod, long-horizon competition under Part D, instead of the two-period interaction under the MDDCP. This broader time horizon introduces considerations of market growth. Thus, prescription drug benefit providers will set prices for a broader horizon, probably considering the trade-off between charging lower prices to attract newcomers and higher prices to already committed consumers. The differences between the two programs notwithstanding, the evidence from the MDDCP does not straightforwardly suggest a secular decline in the level and the dispersion of prices under Part D.

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Federal Reserve Bank of St. Louis *Review*, Annual Index, 2008

JANUARY/FEBRUARY

William Poole, "Thinking Like a Central Banker."

Rajdeep Sengupta and Craig P. Aubuchon, "The Microfinance Revolution: An Overview."

Daniel J. McDonald and Daniel L. Thornton, "A Primer on the Mortgage Market and Mortgage Finance."

Riccardo DiCecio, Kristie M. Engemann, Michael T. Owyang, and Christopher H. Wheeler, "Changing Trends in the Labor Force: A Survey."

MARCH/APRIL

"In Memoriam: Anatol 'Ted' Balbach, 1927-2007."

William Poole, "Market Bailouts and the 'Fed Put.'"

Thomas A. Garrett, "Pandemic Economics: The 1918 Influenza and Its Modern-Day Implications."

Edward Nelson, "Friedman and Taylor on Monetary Policy Rules: A Comparison."

Kristie M. Engemann, Michael T. Owyang, and Sarah Zubairy, "A Primer on the Empirical Identification of Government Spending Shocks."

MAY/JUNE, PART 1

David C. Wheelock, "The Federal Response to Home Mortgage Distress: Lessons from the Great Depression."

William T. Gavin and Geentanjali Pande, "FOMC Consensus Forecasts."

Patrick A. Pintus, "Laffer Traps and Monetary Policy."

William T. Gavin and Kevin L. Kliesen, "Forecasting Inflation and Output: Comparing Data-Rich Models with Simple Rules."

MAY/JUNE, PART 2

Charles T. Carlstrom and Timothy S. Fuerst, "Inertial Taylor Rules: The Benefit of Signaling Future Policy."

Mark A. Wynne, "Core Inflation: A Review of Some Conceptual Issues."

Joseph E. Gagnon, "Inflation Regimes and Inflation Expectations."

Song Han and Casey B. Mulligan, "Inflation and the Size of Government."

JULY/AUGUST

Monetary Policy Under Uncertainty

Lars E.O. Svensson and Noah Williams, "Optimal Monetary Policy Under Uncertainty: A Markov Jump-Linear-Quadratic Approach."

Athanasios Orphanides and Volker Wieland, "Economic Projections and Rules of Thumb for Monetary Policy."

Marek Jarociński and Frank R. Smets, "House Prices and the Stance of Monetary Policy."

James D. Hamilton, "Assessing Monetary Policy Effects Using Daily Federal Funds Futures Contracts."

John B. Taylor, Ben S. Bernanke, and William Poole, "Panel Discussion."

Carl E. Walsh, "Announcements and the Role of Policy Guidance."

William Poole, "Rules-of-Thumb for Guiding Monetary Policy."

Commentaries by:

Timothy W. Cogley
Andrew T. Levin
Charles I. Plosser
Patrick Minford
Robert G. King
Stephen G. Cecchetti
Alec Chrystal
Kenneth N. Kuttner
Marvin Goodfriend

SEPTEMBER/OCTOBER

Michael D. Bordo and Anna J. Schwartz, "Monetary Economic Research at the St. Louis Fed During Ted Balbach's Tenure as Research Director."

Kevin L. Kliesen, "Oil and the U.S. Macroeconomy: An Update and a Simple Forecasting Exercise."

Alistair Milne and Geoffrey Wood, "Banking Crisis Solutions Old and New."

Paul Mizen, "The Credit Crunch of 2007-2008: A Discussion of the Background, Market Reactions, and Policy Responses."

NOVEMBER/DECEMBER

David C. Wheelock, "Changing the Rules: State Mortgage Foreclosure Moratoria During the Great Depression."

Matthew S. Chambers, Carlos Garriga, and Don Schlagenhauf, "Mortgage Innovation, Mortgage Choice, and Housing Decisions."

Christopher J. Neely and David E. Rapach, "Real Interest Rate Persistence: Evidence and Implications."

Emin M. Dinlersoz, Rubén Hernández-Murillo, Han Li, and Roger Sherman, "Drug Prices Under the Medicare Drug Discount Card Program."

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