Mutual Funds and the Evolving Long-Run Effects of Stock Wealth on U.S. Consumption

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Abstract

Lower mutual fund loads have plausibly boosted the stock wealth elasticity of U.S. consumption by enhancing stock liquidity and arguably by inducing stock ownership among middle-income families, consistent with theory and cross-section data (Guiso, Haliassios, and Jappelli (2003), Haliassios (2002), Heaton and Lucas (1996, 2000), and Vissing-Jorgensen (2002)). In load-modified models, the stock wealth elasticity is declining in loads and more stable long-run wealth and income coefficients arise, especially controlling for mortgage refinancing and equity withdrawal activity. Modified models imply that the stock wealth elasticity has risen, while conventional models overestimate the wealth and underestimate the income elasticities of consumption.

JEL Classification Codes: *E21*, *E44*, *G11*

Key Words: Wealth effects, Consumption, and Mutual funds

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A standard explanation for the fall of the U.S. personal saving rate in the late 1990s is that higher stock prices boosted consumption via wealth effects (e.g., Board of Governors of the Federal Reserve System, 2000). Nevertheless, the historical link between consumption and wealth has not been very tight, especially after the saving rate stayed low after the stock market bust of the early 2000s, suggesting that stock wealth effects have either been poorly estimated (Ludvigson and Steindel, 1999) or offset by record levels of families refinancing mortgages or withdrawing equity from housing wealth (Canner, Dynan, and Passmore, 2002). The current paper analyzes whether the time series link between wealth and consumption has been altered by financial innovations that have raised stock and housing wealth liquidity, via lowering mutual fund costs and increasing the proclivity of families to tap housing wealth or refinance mortgages.

Stock wealth enters many econometric models of consumption based on the permanent income and life cycle hypotheses [LCH, e.g., Bosworth (1975), Brayton and Tinsley (1996), and Mishkin (1977)] and on buffer stock approaches. An example of the latter is Carroll (1992), in which households maintain a target wealth-to-income ratio by altering saving, consistent with the 1990's drop in the personal savings rate and jump in the wealth-to-income ratio (Board of Governors of the Federal Reserve System, 2000). Estimated wealth coefficients are often seen as reflecting the impact of underlying economic variables on consumption through their effect on stock prices, whether from altering earnings fundamentals, the risk-less real interest rate, or the equity risk premium. Because the theoretical link between wealth and consumption is strongest over long-run horizons, swings in wealth are generally viewed as having long-term effects.

Nevertheless, the size and stability of the long-run wealth elasticity of consumption have been controversial for at least three reasons. First, some argue that stock wealth has little effect on consumption after controlling for consumer confidence (Hymans, 1970). More recently, Otoo

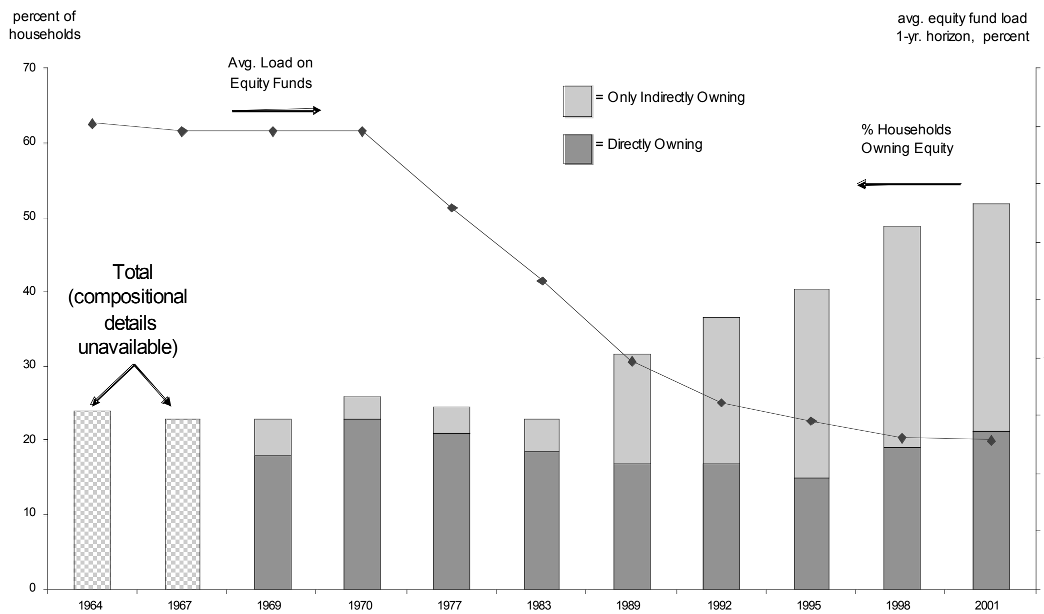
(1999) found that stock prices affect the confidence of stock- and non-stock-owners similarly, suggesting that the information content of stock prices might arise from expectations. Second, estimated stock wealth effects may be unreliable, as in Ludvigson and Steindel (1999), who find that long-run wealth coefficients shift much in rolling regressions. Third, the high concentration of stock assets might limit the size of the long-run wealth elasticity of consumption, as implied by evidence that the rich save differently [Carroll (2000b) and Dynan, Skinner, and Zeldes (2004)] and by models in which labor and capital income uncertainty imply that consumption is concave in wealth under several standard utility functions (Carroll and Kimball, 1996). Alternatively, if the bequest motive is stronger among the wealthy (Carroll, 2000b), the marginal propensity to consume out of wealth is declining in wealth. In either case, the distribution of wealth and income has implications for aggregate consumption (Carroll, 2000a).

My study addresses the second and third criticisms by analyzing whether mutual fund costs add information about the long-run relationship between consumption, income, and wealth. Lower mutual fund costs may raise the aggregate wealth elasticity in two plausible ways. First, lower loads reduce the transfer costs of tapping equity wealth to fund consumption, thereby making consumption more sensitive to stock wealth (Davis and Norman, 1990). Second, because mutual funds were the most feasible vehicle through which middle-class families could own a diversifiable stock portfolio over the sample examined, lower loads in the presence of non-diversifiable labor income risk can theoretically boost stock ownership rates by reducing transfer and entry costs as in Haliassos and Bertaut (1995), Heaton and Lucas (1996, 1997, 2000), and Vissing-Jorgensen (2002).[1] Indeed, equity mutual fund loads at a one-year horizon

[1] A consensus is emerging that, "fixed costs…represent the key factor in stock market participation," (Haliassios, 2002, p. 29), and mutual fund costs are a key component of such costs (Haliassios (2002, p. 29) and differences in mutual fund use help account for cross country patterns of stock ownership (Guiso, Haliassios, and Jappelli, 2003).

have a strong negative correlation (-0.87) with stock ownership rates (Figure 1).[2] If utility is concave in wealth, as Carroll and Kimball (1996) show under plausible assumptions, then wider stock ownership may plausibly boost the aggregate stock wealth elasticity of consumption.

Figure 1: Equity Fund Loads Fall and Stock Ownership Rates Rise



This study assesses whether accounting for stock fund loads can improve estimates of the long-run relationship between consumption and stock wealth. Specifically, the product of loads and stock wealth is added to models, where this term should be negative if falling loads boost stock liquidity or equity participation and thereby, the long-run sensitivity of consumption to

[2]Stock ownership data are not *fully* consistent for some minor definitional differences across SCF's (see Duca (forthcoming)). The impact of these differences appears minor—for details, see an appendix available upon request from the author. Ownership data are from Aizcorbe, et al. (2003), Durkin and Elliehausen (1978), Katona et al. (1968, 1970, 1971), and Kennickell, et al. (2000). SCF data on stock wealth distribution are likely less consistent than ownership rates. The former are affected by factors altering the consistency of ownership rates and by small samples of the rich in early SCFs that gave rise to large discrepancies between SCF and IRS data. Since 1983, SCF have over-sampled the rich. Owing to their dichotomous nature, ownership rates are less affected by such problems and loads plausibly affect ownership rates more than distribution since the rich can diversify without mutual funds.

stock wealth.[3] This study focuses on empirical results to address some of the controversy over stock wealth effects and, for theoretical motivation, refers to papers on equity participation, the concavity of consumption, and the impact of transfer costs on portfolios.

Because the sample covers the early 2000s, this study assesses how consumption was affected by greater mortgage-related activity in the form of large swings in mortgage refinancing volume and mortgage equity withdrawals (Canner, Dynan, and Passmore, 2002). This mortgage activity marks a divergence from earlier behavior in two ways. First, estimates provided below indicate that aggregate mortgage interest saving from refinancing has recently surged, reflecting plunging mortgage rates and a greater tendency for households to refinance for a given amount of interest rate saving (Duca and Anderson, 2004). Second, applying an approach similar to that of the Bank of England (Davies, 2002) to U.S. Flow of Funds data indicates that mortgage equity withdrawals have accelerated from levels noted earlier by Canner, Dynan, and Passmore (2002). Because these two changes in mortgage-related behavior may have substantial effects on consumption during a period of shifting stock wealth, controlling for them in some specifications serves as an important robustness check against possible misspecification and as a means of evaluating competing explanations for shifts in long-run income and wealth elasticities.

This paper is organized along the following lines. As a prelude to the empirical results, the next section discusses how transaction costs can theoretically affect the magnitude of the long-run stock wealth effect on consumption. Section 3 presents the econometric models and data, with details on mutual fund and mortgage activity variables. Section 4 reports the empirical results, which are summarized and interpreted in the conclusion.

[3] This interactive term may also track changes in stock liquidity. While direct costs of buying stocks have fallen by about 50 basis points since 1980 (Jones, 2000), one-year loads have fallen by a larger 500 basis points, consistent with the post-1980 rise in overall equity ownership rates stemming from greater indirect ownership. This study focuses on stock-ownership rates and liquidity, rather than home-ownership rates, as discussed in the appendix.

2. Why Lower Transactions Costs Could Theoretically Increase the Wealth Effect

There are at least two basic, theoretical ways in which lower transactions costs could increase the size of the long-run effect of stock wealth on consumption.

Enhancing the Liquidity of Stock Wealth

First, by cutting the cost of accessing wealth, lower transactions costs raise the propensity to consume out of wealth. It is theoretically plausible that agents are more likely to realign portfolios if they face lower costs when transferring assets. Liu and Loewenstein (2002) and Zakamouline (2002) find that transactions costs create a zone where it is optimal not to trade between risky and safe assets until portfolio misalignment becomes large enough to justify the transactions cost of realigning assets. As proportional transfer costs rise, this "no trading" zone widens. Davis and Norman (1990) find similar results assuming that utility is characterized by hyperbolic absolute risk aversion in a world of two assets: stocks having variable returns of a known variation and bank deposits yielding fixed returns. As proportional asset transfer costs fall, the smaller is the range of stock price variation under which agents do not alter their portfolios and consumption in response to changing equity prices.[4] For middle-income families, diversification and limited wealth imply that mutual fund loads are the relevant transfer cost.

Broadening the Distribution of Stock Wealth

For families with limited wealth, mutual funds were likely the most feasible way to own diversified stock portfolio prior to exchange traded funds, and equity fund loads were arguably the most relevant transfer cost affecting equity participation. Recent models of equity participation imply that lower transactions costs of investing stocks (e.g., mutual fund loads)

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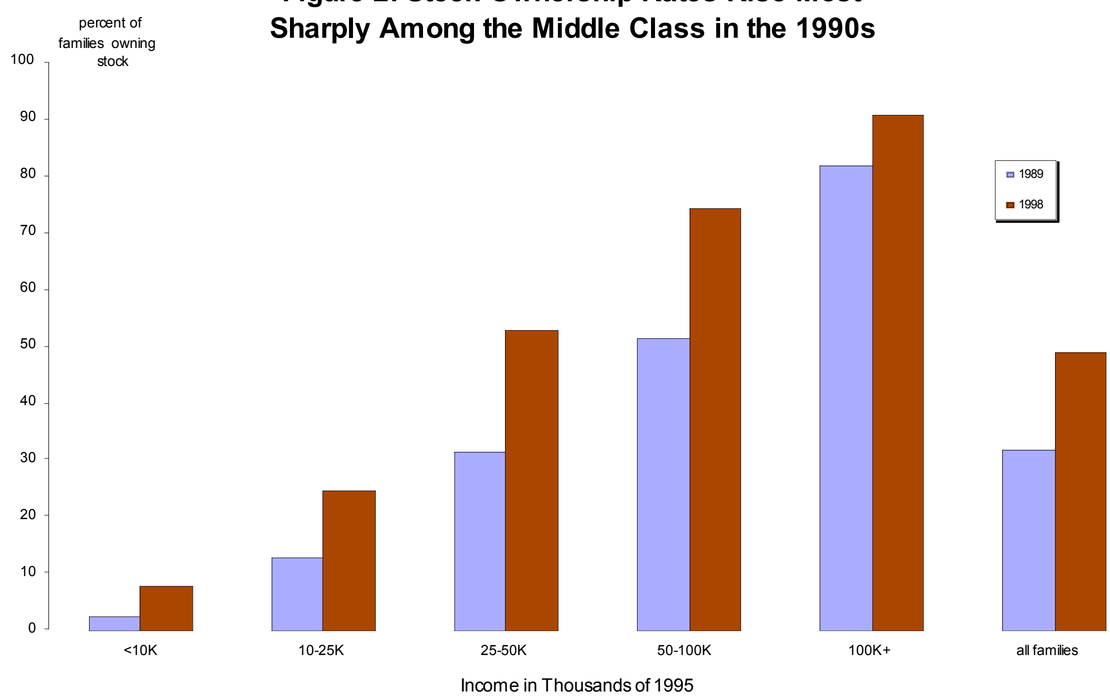
[4]While it is hard to separate liquidity from stock ownership effects in this study, other studies find that markets value liquidity in that stock returns are positively correlated with bid-ask spreads [Amihud and Mendelson (1986a,b), Chalmers and Kadlec (1998), Datar, Naik, and Radcliffe (1998), and Swan and Westerholm (2000)].

could induce higher stock ownership rates (Heaton and Lucas (2000, 1997, 1996) and Vissing-Jorgensen (2002)), particularly for middle-income families (e.g., Heaton and Lucas, 2000).

Several facts are consistent with this view. First, the rise in the equity ownership rate (Figure 1) owed to increased indirect ownership, reflecting a greater relative use of mutual funds as a means of owning stock (Duca 2005). Second, increases in stock ownership rates were largest among middle-income households, for whom the diversification appeal of mutual funds is relevant (Figure 2). Third, higher stock ownership rates are not largely due to shifts in the age profile of the U.S. population. In fact, Laderman (1997) finds that demographic shifts cannot account for higher mutual fund ownership through the mid-1990s, which instead reflected

Figure 2: Stock Ownership Rates Rise Most Sharply Among the Middle Class in the 1990s



increases in ownership rates within most age cohorts. Rising stock ownership seems to have been induced by lower equity fund loads stemming from improved financial productivity, as implied by Duca's (2005) findings that equity fund loads are cointegrated with the rising use of equity funds, equity fund loads are cointegrated with banking sector productivity, and that improvements in financial sector productivity slightly led declines in mutual fund loads, which in turn slightly led increases in household use of mutual funds to own equity. Thus, higher equity participation likely owes to technology than to demographic shifts or a general rise in stock market efficiency (which would have equally boosted direct and indirect stock ownership).

Greater equity participation induced by lower equity fund loads could boost the overall stock wealth effect if the propensity to consume out of wealth among low and middle-class households exceeds that of the wealthy. Consistent with the findings of Dynan, Skinner, and Zeldes (2004), if bequests are increasing in wealth or the very wealthy have stronger saving preferences, the ratio of consumption to stock wealth is increasing in equity participation. With respect to the former possibility, Carroll (2000b) shows that under CRRA preferences [$u(c_t) = \{c^{1-\rho}\}/(1-\rho)$], if bequests yield utility in a Stone-Geary form [$v(w) = \{(w+\gamma)^{1-\alpha}/(1-\alpha)\}$], then bequests are a luxury good if the rate of time preference (ρ) exceeds (α), implying the propensity to consume out of wealth is declining in wealth. With respect to the latter possibility, Gentry and Hubbard (1998) find that people with high risk preferences who tend to earn higher returns, tend to save more to self-finance entrepreneurial investment owing to capital market imperfections.

Theoretically, larger stock wealth effects could also be linked to broader stock ownership if uncertainty affects utility. Carroll and Kimball (1996) show that a precautionary savings motive arises ($u'''u'/u''^2 = k>0$) and consumption is concave in wealth under reasonable conditions if utility exhibits hyperbolic absolute risk aversion, which encompasses constant relative and constant absolute risk aversion. If risks to future capital returns and labor income are imperfectly correlated, consumption is strictly concave in wealth. This implies that as stock ownership rates rise, the sensitivity of aggregate consumption to wealth should increase.

3. Empirical Specification and Data

The analysis primarily uses the dynamic ordinary least squares (DOLS) approach of Stock and Watson (1993) to estimate long-run relationships among consumption, income, and wealth. This approach differs from the older Johansen-Juselius approach by taking into account information contained in leads of first differences of long-term variables. In properly implementing the DOLS approach, the lag length on leads and lags of first differenced variables needs to be long enough to ensure that residuals are stationary. The estimated models include a standard model, a mutual fund-modified model, and versions that include mortgage variables.

The Standard or Baseline Model

The baseline or standard model of consumption is of the form:

$$c_t = \alpha + \beta_0 y_t + \delta_0 w^s{}_t + \gamma_0 w^o{}_t + \Sigma_{i=[1,3],[-3,-1]}[\beta_i \Delta y_{t+i} + \delta_i \Delta w^s{}_{t+i} + \gamma_i \Delta w^o{}_{t+i}], \qquad (1)$$

where lower case letters denote logs, y = labor plus transfer income (omits property and wealth income), w^S = stock wealth, and w^O = other wealth. Two measures of consumption are used. One is spending on consumer non-durable goods and services (C^{ns}), which excludes durable purchases because durable consumption equals the service flow from the stock of durables. As in the Federal Reserve Board's econometric model, the second measure (C^{nds}) adds the service flow from the durable stock to C^{ns}. Three leads and lags of first difference terms are sufficient to generate stationary errors. To avoid complications surrounding rationing during the Korean War and possibly large revisions to very recent data estimates, regressions are estimated using data spanning 1954:1-2002:4. Because 3 leads and lags are used, the availability of load data since 1959:Q4 permits equations (1) and (2) to be estimated over the sample 1955:1 to 2002:1.

Eq. (1) is similar to the model of Ludvigson and Steindel (1999), but differs in two ways. First, because transfer income often offsets swings in labor income, the income term includes transfer income, which improves model fit. Second, to test whether equity fund loads affect stock wealth coefficients, eq. (1) departs from Ludvigson and Steindel by separating wealth into stock and other assets. This is done because it is unlikely that mutual fund costs track the transaction costs of non-stock assets, such as real estate fees. In addition, the main hypothesis tested is whether stock ownership rates affect the stock wealth channel and it is the link between equity participation and loads that mainly justifies using loads to proxy for distribution effects.

The Mutual Fund-Modified Model

The modified model adds the product (w^sL) of log stock wealth (w^s) and the level of the average equity fund load (L), and adds 3 leads and lags of its first difference to eq. (1):

$$c_t = \alpha + \beta_0 y_t + \delta_0 w^s_t + \phi_0(w^s_t L_t) + \gamma_0 w^o_t + \Sigma_{i=[1,3],[-3,-1]}[\beta_i \Delta y_{t+i} + \delta_i \Delta w^s_{t+i} + \phi_i \Delta(w^s_{t+i} L_{t+i}) + \gamma_i \Delta w^o_{t+i}] \quad (2)$$

Eq. (2) allows equity fund loads (L) to affect the elasticity of consumption with respect to stock wealth. In levels, this is equivalent to treating long-run consumption (C*) as

$$C^* = \alpha(Y)^{\beta}(W^s)^{(\delta+\phi L)}(W^o)^{\gamma}, \quad (3)$$

where it is hypothesized that $\beta>0$, $\gamma>0$, $\delta>0$, and $\phi<0$. One qualification about using L to proxy for stock ownership rates is that doing so runs the risk of measuring an explanatory variable with error, which can potentially result in serious estimation problems.

While the income and wealth terms are conventional, the equity mutual fund load variable is not. Following Duca's (2000, 2005) studies of how mutual fund costs have affected M2 balances and household equity portfolios, respectively, loads are measured using fees from a sample of large equity mutual funds that are aggregated into an average based on each fund's relative share of assets under management in the sample. Details on Duca's panel are as follows. One measure, L, tracks near-term loads using the weighted-average front-end and back-end load (as a percentage of asset transfers), where the back-end load is for withdrawals within a year of investment, and fund weights are the ratios of each fund's assets to all funds sampled. Since L is assumes investors have a short horizon, an alternative, L5, employs a five-year horizon, using front-end loads divided by five and back-end loads for withdrawals after five years, also divided by five. In regressions not shown, results are similar using L5 in interactive stock wealth terms.

While L and L5 may be distorted if mutual funds raise expense ratios when cutting loads, results are similar if expense ratios are added to L and L5, in line with similar trends in L, L5, and annual (post-1980) estimates of Rea and Reid (1998), who show that total equity fund costs fell in the 1990s as lower loads outweighed slightly higher expense ratios. Better technology likely cut costs as implied by Duca's (2005) findings that the broadest measure of equity fund costs (L5 plus expense ratios) is cointegrated and negatively related to productivity in banking (the only financial sector for which productivity data are available), and that long-run movements in bank productivity embodied in an error-correction term lead short-run changes in mutual fund costs (and portfolio behavior), with no reverse causality.

Because data before the mid-1980s are sketchy and incomplete, mutual fund costs were based on a sample of large mutual funds. Funds were selected if their assets were at least $1 billion at year-end 1991 if the fund existed before the mid-1980s; were at least $2 billion at year-end 1994 if the fund's inception date occurred after 1983; were at least $5 billion at year-end 2003; or were at least $250 million at year-end 1975. The first criterion reflects whether a fund was sizable during the early 1990s, when bond and equity mutual fund loads fell and households shifted assets out of M2 into mutual funds. The second criterion reflects whether a growing but new fund was large near the end of these portfolio shifts out of M2 (Duca, 2000). The third criterion reflects whether a fund remained large following the stock market bust of the early 2000s. Given the stock appreciation of the 1990s, the hurdles for newer funds were higher for the 1994 and 2003 cutoff dates to keep data gathering costs from exploding. The fourth criterion avoids excluding funds that were relatively large in 1975 from distorting averages when fewer funds existed. Also excluded were funds that were closed-end, closed to new investors, only open to employees of a specific firm, or institutional. One member, the Windsor Fund, became closed to new investors, but was included because its cousin (Windsor II) was started when the Windsor fund became closed to new investors and both funds are large. Also included because of its size is the Magellan Fund which became closed to new investors in 1998. 133 equity funds are in the sample (a list is available) using data from the funds and various issues of Morningstar, IBC/Donoghue, and CDA/Wiesenberger (a, b).

L and L5 use quarterly loads of sampled funds. Since only year-end asset data on mutual funds are available, quarterly asset weights are interpolated from a year-end data (annual weights are commonly used to construct quarterly data) and the quarterly inception dates of funds. Since year-to-year changes in asset weights are small, using interpolated quarterly weights poses few problems. In contrast, annual data indicate a much larger impact of changes in fund load fees on annual changes in weighted-average loads. Since quarterly data on individual fund load fees are combined with interpolated quarterly asset weights, the series likely tracks asset-weighted quarterly load changes well. This also poses little handicap because the study focuses on the long-run relationship between stock wealth and consumption and since most quarterly changes in the interactive w^sL variable reflect swings in stock wealth (w^s). Standard tests indicate unit roots for L, w^sL and the log levels of the income, consumption, and wealth variables used.

Mortgage Activity Modified Models

To assess the impact of changing mortgage activity on consumption, I(1) variables tracking mortgage refinancing activity (REFI) and mortgage equity withdrawals (MEW) were added to the baseline and load-modified models (equations (1) and (2), respectively), along with three leads and lags of their first differences. Because some observations are zero or negative, these variables are not in logs, but are constructed as ratios to non-property income.[5]

Following Duca (1990) and Anderson (1993), quarterly interest cost savings from mortgage refinancings were constructed using three data series. First, the dollar volume of mortgage refinancing was tracked using Anderson's data on mortgage-backed security (MBS) liquidations by Ginnie Mae, Fannie Mae, and Freddie Mac (Duca and Anderson, 2004). This approach has the advantage of tracking dollar volumes since 1970, in contrast to the Mortgage Bankers' Association (MBA) index of applications for mortgage refinancing, which indexes the number—not the dollar volume—of refinancing applications since the early 1990s.[6] The approach used here allows the interest rate savings from refinancings to be measured separately from mortgage equity withdrawals, which stem from home equity borrowing, home sales, and cash-out refinancings. The main drawback of the MBS approach is that it omits the refinancing of mortgages that were not packaged into MBS's. However, the MBS series may consistently track the aggregate volume of refinancing activity since Duca and Anderson (2004) find that the MBA index and MBS liquidations line up closely over a common sample. To convert refinancing volume into interest cost saving, the average reduction in interest rates is measured using the ratio of interest rates on the old to the refinanced mortgage multiplied by the effective rate on new 30-year fixed-rate, conventional mortgages (Freddie Mac). This rate gap is multiplied by the amount of refinanced MBSs to calculate dollar savings from refinancing (real and annualized). Next, since these mortgage cost savings can aid discretionary spending for a

[5] The greater propensity for households to refinance mortgages and tap mortgage equity wealth reflects several factors, such as technology (lower fixed costs and less time), tax changes, and home sales (which affect equity withdrawals). For purposes of measuring income and wealth elasticities, the two mortgage variables implicitly control for these factors. The cointegrating vectors that have mortgage variables are already large and it becomes more difficult to identify unique cointegrating vectors as the number of variables rises. Each additional variable also uses up seven degrees of freedom in DOLS regressions, posing limits on the usefulness of rolling regressions.

few years, the interest savings of mortgage refinancings over the past three years is cumulated and divided by non-property income to create REFI. This ratio to non-property income is used because refinancing savings are zero in many quarters and the consumption models are in logs.

Similar to the Bank of England's series for the U.K. (Davey, 2001), mortgage equity withdrawals (MEW) equal net mortgage borrowing (including home equity loans) minus household residential investment (seasonally adjusted, Flow of Funds data, table F100) divided by non-property income. The series differs slightly from the Bank of England's in not netting out the costs of selling homes or originating mortgages, which are unavailable over the long sample.[7] In most years of the sample, households tended to invest more in housing than they borrowed, partly reflecting the tendency for new home buyers to provide down-payments. However, this has reversed in recent years, when mortgage equity withdrawals jumped to as much as 5 to 6 percent of non-property income at a time of substantial refinancing activity (Figure 3). Nevertheless, mortgage equity withdrawals and the interest reduction from mortgage refinancing activity have sometimes diverged, such as in the mid-1990s when mortgage refinancing jumped and mortgage equity withdrawals fell back. For this reason, it may be empirically important to control for both types of activity by including both REFI and MEW.

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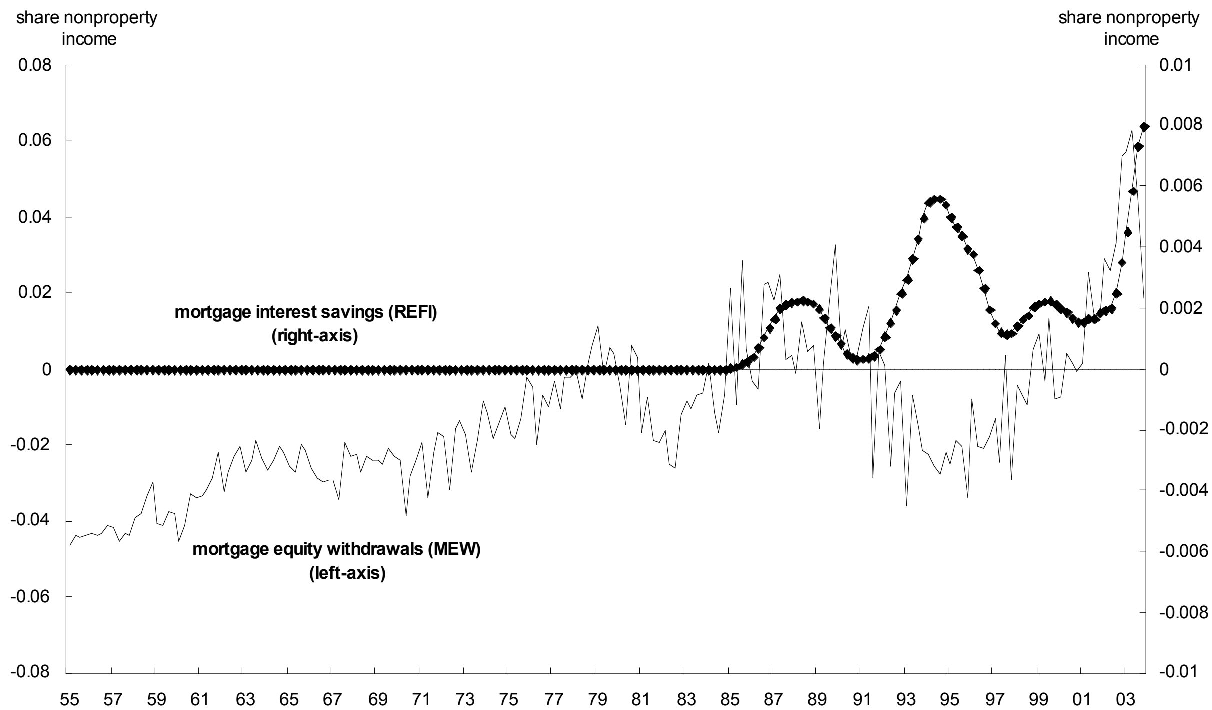
[6] Note that REFI was set to 0 before 1970, which is reasonable given the rising trend of mortgage rates from the mid-1950s to 1970 and the low level of refinancing activity from 1970 to the mid-1980s.

[7] U.S. existing home sales data begin in 1966 and data on average selling costs per home are unavailable (standard realtor commissions do not apply to sales directly made by owners). Mortgage cost data (including points, Freddie Mac) begin in 1970 and vary enough to make it hazardous to set them at some fixed average before 1970.

Figure 3: Mortgage Equity Withdrawals and Cost Savings From Mortgage Refinancings Recently Surge, But Have Not Always Moved in Lock-Step



4. Empirical Results

Results from full sample Johansen and DOLS models are presented. Then rolling regression DOLS results are reviewed to assess the stability of the elasticity estimates. Finally, DOLS results are reevaluated adding a time trend interacted with stock wealth to the models.

4.1 Full-Sample Cointegration and DOLS Results

Johansen cointegration tests are discussed (Table 1) before the DOLS results, with lag lengths set at the minimum needed to find a unique cointegrating vector using the rank criterion (trace statistic) allowing for a deterministic time trend in each variable (which is reasonable for income and consumption), but not a time trend in the vector. A unique cointegrating vector can be found in each case, with non-property income (y), non-stock wealth (w^o), and stock wealth having positive and statistically significant coefficients (cointegrating vector signs are reversed to express equilibrium consumption as a function of variables). One difference is that the income coefficients are higher and more reasonable in the load-modified models, where the trace statistic is more significant using the broader concept of consumption.[8] In the load modified models (columns 2, 4, 6, and 8), the interactive term (w^sL) is significant and negative. This implies that higher loads are associated with a smaller, overall positive impact of stock wealth on consumption, equal to the coefficient on w^s plus the product of L and the coefficient on w^sL.)

The refinancing variable has a positive and significant effect that is larger in the non-load modified models, but the mortgage equity withdrawal variable (MEW) has the anticipated positive and at least marginally significant coefficient only in the load modified models. These and other aspects of the mortgage variables are discussed further with the DOLS results. The DOLS approach to estimating cointegration is plausibly superior to the Johansen-Juselius approach in estimating consumption with current income data because consumption theoretically depends on permanent or expected income. As Ludvigson and Steindel (1999) note, the reason is that the DOLS approach implicitly tracks permanent income better because unlike the Johansen-Juselius approach which includes lags of first differences when estimating cointegrating vectors, the DOLS approach also includes leads of first differences of variables when estimating long-run coefficients, thereby minimizing bias in estimating long-run income effects if consumption anticipates future income, consistent with most theories of consumption.

DOLS models are estimated using equations (1) and (2). As shown in the odd-numbered columns of Table 2, the standard model yields reasonably signed coefficients. However, the corresponding load modified models (columns 2, 4, 6, and 8, respectively) yield more plausible

[8] In this sense, the Johansen tests for load-modified models favor using a broad measure of consumption, consistent with the notion that cointegration tests involving income and spending should use broader concepts of both owing to

magnitudes of the income elasticity of consumption. Also, the stock and non-stock wealth coefficients are larger in the corresponding standard models, which imply a fixed stock wealth elasticity of 3.7 to 5 percent for narrow consumption (c^{ns}) and 5 to 6-½ percent for broader consumption. Consistent with the hypothesis regarding loads and wealth, the coefficient on the interactive term w^sL is negative and statistically significant in each corresponding load-model. This finding, coupled with a smaller non-interactive stock wealth coefficient in corresponding load-modified models implies that the overall stock wealth elasticity is smaller in load models than in corresponding standard models because the overall stock wealth elasticity in each load-modified models equals the coefficient on w^s plus the product of L and the coefficient on w^sL.

Thus, the relatively high loads of the 1960s and 1970s imply a lower overall stock market wealth effect in the early part of the sample and the lower loads of the late part of the sample imply a larger overall effect in recent years. This is illustrated in Figure 4, which plots the sum of the coefficient on stock wealth with the product of L and the coefficient on w^sL from the load and mortgage modified model of narrow consumption (fourth column) with the stock elasticity from the corresponding standard model (third column). Notice that although the overall stock-wealth elasticity has risen according to the load-modified model, it is always below the stock-wealth elasticity estimate of the baseline, fixed-effects specification. Related to this pattern—common to each comparison of corresponding standard and load-modified models—is that the load-modified models yield larger income coefficients in the full sample. This may arise because the load models are arguably better specified, as suggested by rolling regressions yielding stock and income coefficients that vary less over time. In particular, the larger income and smaller wealth coefficients may reflect that load modified models put less weight on stock wealth early in the sample when stock wealth and income are more closely correlated.

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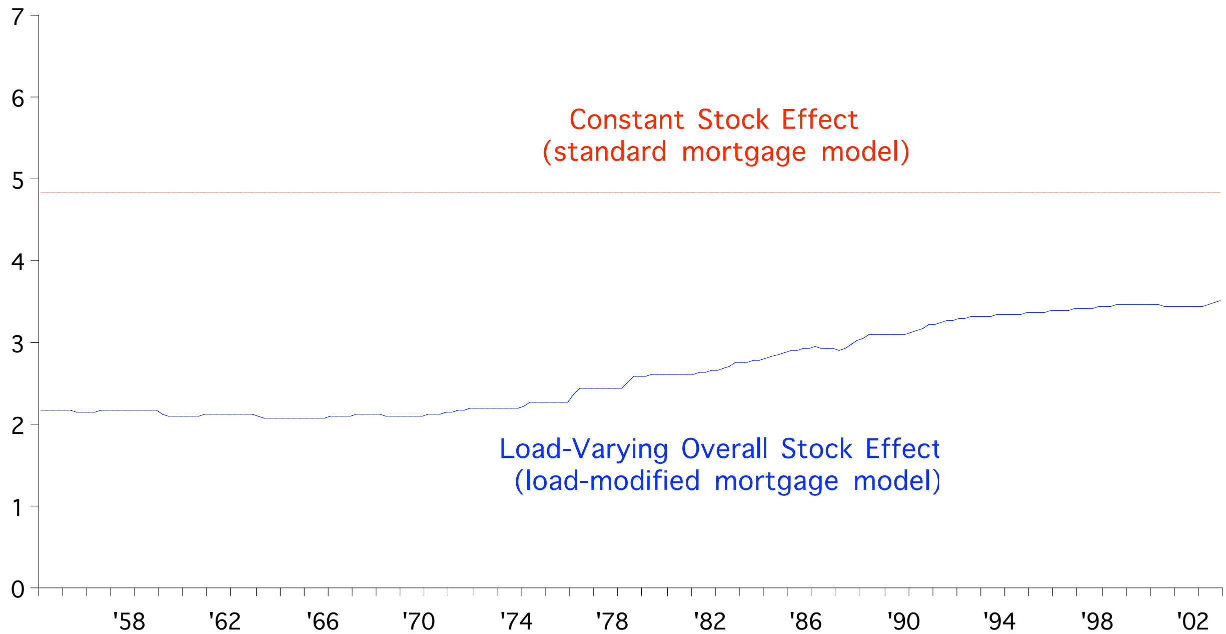
the broad nature of the long-run household budget constraint (see Rudd and Whelan, 2002, for more on this issue).

Figure 4: Stock Wealth Elasticity of Broad Consumption (Mortgage



Results from including mortgage activity variables have similar and dissimilar aspects across the corresponding models. As in the Johansen estimates, the refinancing variable is positive, above unity, and significant in each specification (3, 4, 7, and 8), with larger coefficients in the corresponding standard models. For two reasons, the long-run elasticity of consumption with respect to REFI could exceed unity. First, households may spend more than the interest savings if they re-lever their balance sheets, perhaps even restoring their previous debt service burden (McKelvey, 1994). Indeed, despite record refinancing activity, Federal Reserve data show only a modest dip in household debt service (financial obligations) burdens in the early 2000s in contrast to the sharper pullback during the slow economy of the early 1990s.

Given that recent debt-payments-to-income ratios are around 14 percent, a full re-levering of mortgage interest cost savings to finance spending would boost spending by a factor near seven (roughly the inverse of the debt burden ratio). Second, the variable does not track all interest rate savings, since REFI omits cost savings from the refinancing of mortgages that were not packaged into MBS's. In fact, roughly half of all household mortgage debt was packaged into MBS's in recent years according to Flow of Funds data. Combining these factors, the seemingly large coefficients on REFI in Tables 2 and 3 cannot be summarily dismissed as being too large.

Compared to findings regarding refinancing effects, results for the mortgage equity withdrawal variable differ not only quantitatively, but also qualitatively when comparing models. MEW is statistically insignificant in the standard stock wealth specifications (3 and 7), but has a positive and highly significant impact in load-modified models (4 and 8).

4.2 Estimates From Rolling DOLS Regressions

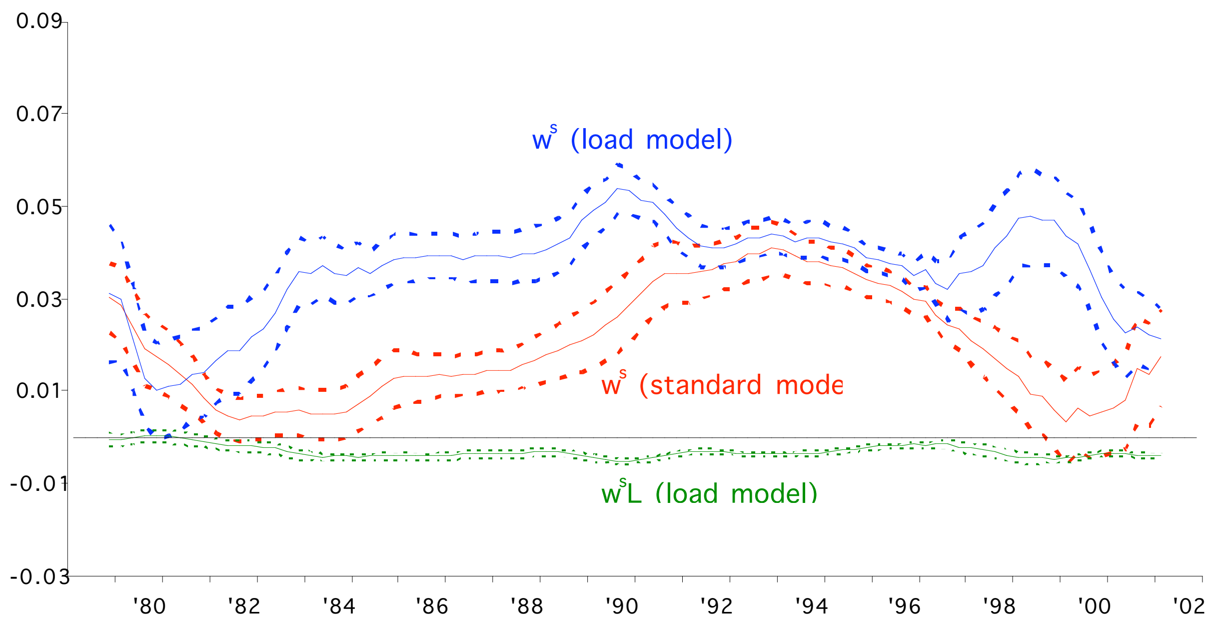
Rolling regressions are run using a 25-year horizon to assess the reliability of estimates. Unlike Ludvigson and Steindel (1999), who use a 10-year rolling horizon for their more parsimonious model, a 25-year rolling horizon is used because there are up to four extra sets of wealth and mortgage activity terms (the level and 6 first difference terms for each term) that use up as many as 28 more degrees of freedom. 25 years allows the 43 coefficients of models 4 and 8 to be estimated with 56 degrees of freedom, which is reasonable for estimating long-run effects.

From models 5 and 6, Figure 5 plots the coefficients on w^s and w^sL from the load-modified model, along with 1 standard-error bands.[9] In rolling regressions, the coefficient on w^s from the standard model shifts a great deal and nears zero, in contrast to that from the load-modified model, which moves less over time. This finding is not the by-product of instability in

[9]Broader consumption yielded coefficients that varied less over time, but the same qualitative comparisons between the standard and load-modified models hold when using consumption of services and nondurable goods.

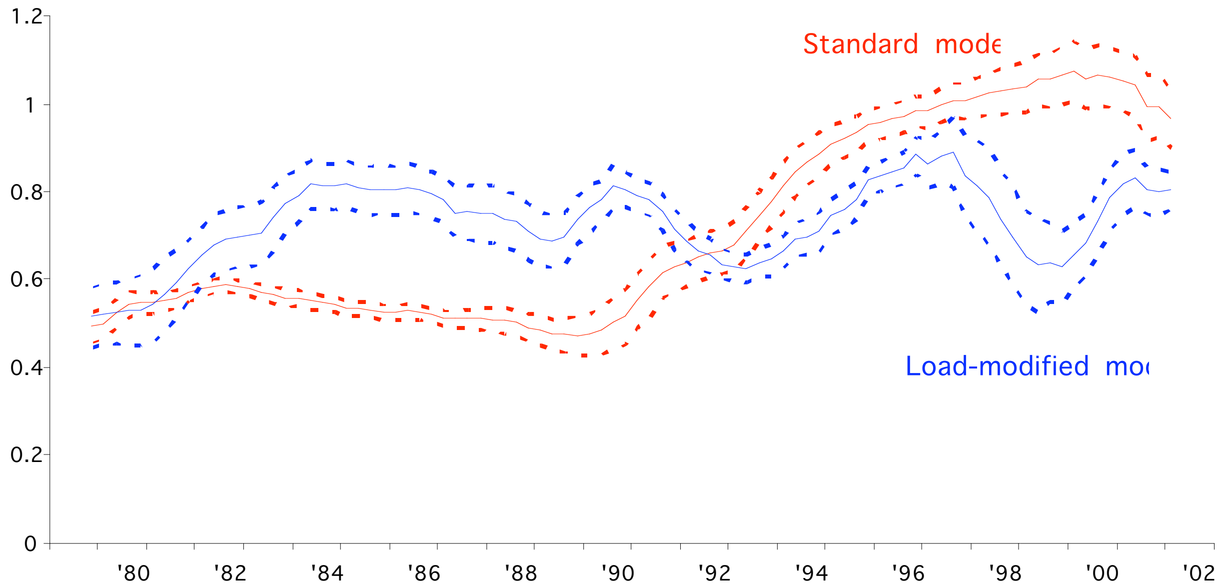
the coefficient on w^sL, nor in the estimated income elasticity, which shifts less over time in the corresponding load-modified models (Figure 6). Although the coefficient on w^sL is about zero in samples ending before the mid-1980s, this largely reflects that loads were flat through most of the 1970s and thus add little information in earlier years. In rolling samples ending after the early 1980s, which include periods when loads differ, the coefficient is negative and relatively stable. These different patterns for stock wealth and income elasticities are more apparent in the mortgage-modified models (see Figures 7 and 8, respectively). Finally, another indication that accounting for mutual funds can improve parameter stability is that the rolling regression coefficients in the load-modified models are near the full-sample DOLS estimates (Table 2) in Figures 5-8, whereas this is not the case for coefficients from the standard (non-load) models.

Fig. 5: Stock Wealth Coefficients from 25-Year Rolling Regress
(Non-Mortgage Models of Broad Consumption)



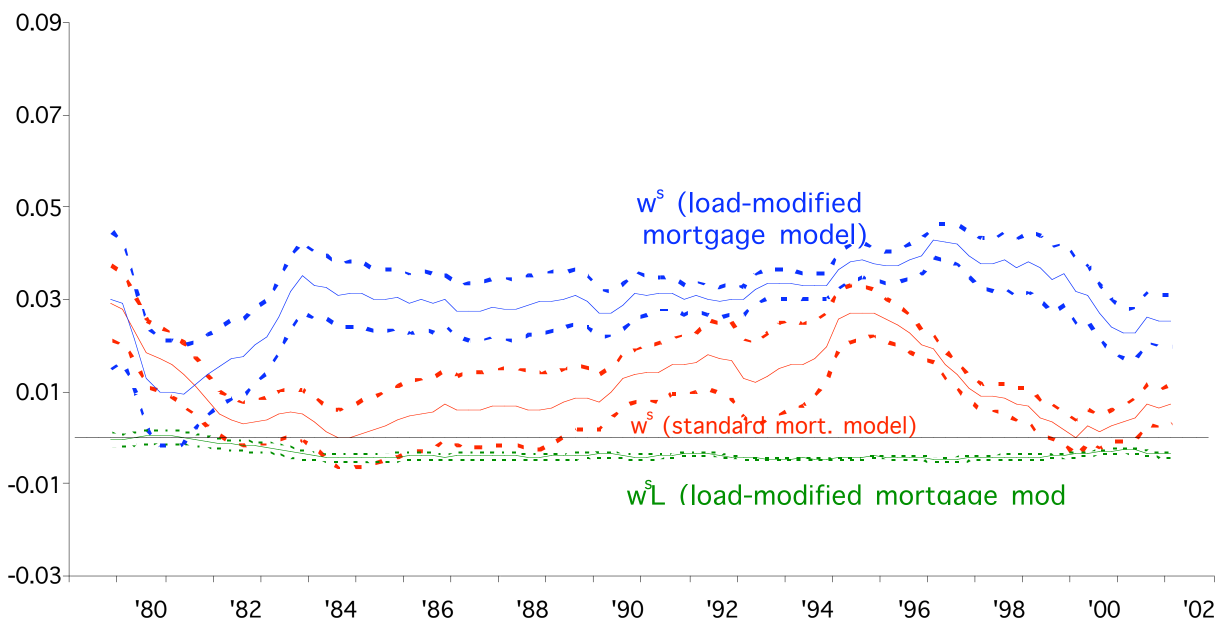
Note: dashed lines denote one-standard deviation bands.

Fig. 6: Income Elasticity Coefficients from 25-Year Rolling Regre
(Non-Mortgage Models of Broad Consumption)



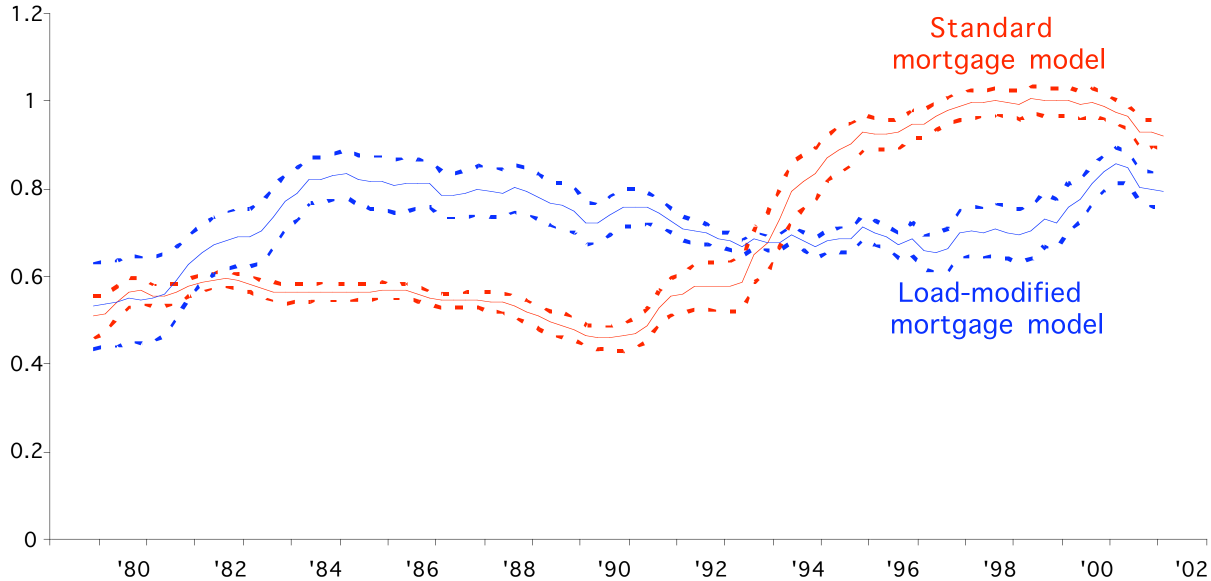
Note: dashed lines denote one-standard deviation bands.

Fig. 7: Stock Wealth Coefficients from 25-Year Rolling Regress
(Mortgage Models of Broad Consumption)



Note: dashed lines denote one-standard deviation bands.

Fig. 8: Income Elasticity Coefficients from 25-Year Rolling Regre
(Mortgage Models of Broad Consumption)



Note: dashed lines denote one-standard deviation bands.

4.3 Load Versus Time Trend Effects on Stock Wealth Elasticities

One concern is that loads may pick up a time trend, even though loads are flat until 1973. To address this issue, the product of a time trend and wealth is added to the models in Table 2.[10] As shown in Table 3, the time-stock wealth term is significant and positive in standard models (columns 1, 3, 5, and 7), indicating a growing sensitivity to stock wealth over time. However, in the load models, the interactive load (w^sL) terms remain at least marginally significant and are more significant than the interactive time (w^s*T) variables (columns 2, 4, 6, and 8). In addition, signs on the time-wealth variable shift from positive in two conventional models (column 1 and 3) to negative in corresponding load models (columns 2 and 4), while among the models of broad consumption, the time-wealth term is insignificant in one load-model (column 6). In only

[10] Leads and lags of the first differences of w^sL are omitted because they are co-linear with those of w^s.

one case is the time-stock wealth term positive and marginally significant in a load-model (column 8), but it is less significant than the interactive load-stock wealth variable.

Two other patterns of results cast doubt on the approach of adding interactive time/wealth variables. First, the non-interactive stock wealth (w^s) terms remain significant in the load models when the time-interactive stock wealth (w^sT) variable is present, but become insignificant in the time-modified models of broad consumption (columns 5 and 7). Second, the coefficients on the non-interactive and time-interactive stock wealth terms from time-modified models move much in rolling regressions, as do the income coefficients (charts and an appendix demonstrating this point are omitted to conserve space, but are available from the author). Thus, using loads outperforms using a time trend in several empirical dimensions, in addition to providing an economic rationale for why long-run stock wealth effects have become more important.

5. Conclusion

Through enhancing the liquidity of stock wealth and by inducing higher stock ownership rates, falling equity mutual fund loads could theoretically increase the aggregate long-run wealth elasticity of consumption. The latter channel partly depends on whether utility is concave in wealth, which can arise if uncertainty affects utility (Carroll and Kimball, 1996) or if the bequest motive is stronger among the very wealthy (Carroll 2002). If utility is concave in wealth and if a decline in the transaction costs of owning a diversified portfolio of stocks induces less wealthy households to become shareholders, then lower transactions costs can plausibly induce an increase in the long-run impact of stock wealth on aggregate consumption. This study tests whether equity fund loads add information about the long-run relationship between income, stock wealth, other wealth, and consumption in the U.S. Findings indicate that the stock wealth elasticity of consumption is declining in equity fund costs and that more stable long-run relationships between consumption and stock wealth and between consumption and income are obtained when accounting for such costs. In this sense, greater market efficiency in the form of lower mutual fund costs has ostensibly boosted stock liquidity and equity participation and accompanied an underlying increase in the long-run elasticity of consumption with respect to stock market wealth in the United States.

These results help reconcile Ludvigson and Steindel's (1999) finding that conventional estimates of long-run stock wealth effects are imprecise with evidence (1) that stock prices affect the spending of shareholders more than non-shareholders (see Dynan and Maki (2001) for U.S. evidence and Attanasio, Banks, and Tanner (2002) for U.K. evidence) and (2) that the personal saving rate and the wealth-to-income ratio are negatively correlated. The findings also accord with microeconomic evidence (e.g., Carroll, 2000a) that shifts in the distribution of wealth can have implications for the behavior of consumption at the aggregate level.

With respect to macro-policy implications, the load-modified models imply that although the long-run wealth elasticity of consumption has risen over time, conventional models appear to overestimate the wealth elasticity and underestimate the income elasticity over long samples, especially when controlling for mortgage refinancing and equity withdrawal activity. These last factors have become increasingly important in light of the greater tendency for households to refinance mortgages and tap housing equity. Overall, findings indicate that it important to account for innovations that have reduced the costs of investing in mutual funds and of tapping housing wealth when analyzing consumption behavior in the United States.

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Appendix: Why This Study Focuses on How Changes in Stock-Ownership and Mortgage Activity, Rather Than Changes in Home-Ownership Rates, Could Affect Consumption

The paper focuses on stock- rather than home-ownership effects for several reasons. First, the homeownership rate fluctuated in a narrow range (63-66%) over most (1965-2002) of the sample (see Figure 9) and has only increased slightly beyond this range in recent years, from 66% to 68%. This is in contrast to the stock ownership rate, which more than doubled in the last 25 years, from around 25% in 1977 to 51.9% in 2001, with most of the increase occurring by 1998, when the equity participation rate reached 48.9%. With the exception of the 2-percentage point increase in homeownership rates since the late 1990s, other data indicate that the post-WW II rise in homeownership rates largely occurred in the late 1940s and 1950s, which predates most of the sample used in estimation.

Third, shifts in the income coefficient from rolling regressions of the conventional model showed up well before homeownership rates increased in the late 1990s. Fourth, while there has certainly been an increase in the liquidity of housing wealth, most of this greater liquidity has occurred only in recent years and the analysis attempts to control for these effects by including estimates of mortgage cost reductions from refinancing activity and of mortgage equity withdrawals.

Figure 9: U.S. Homeownership Rates Stayed Within a Narrow Range Until the Late 1990s

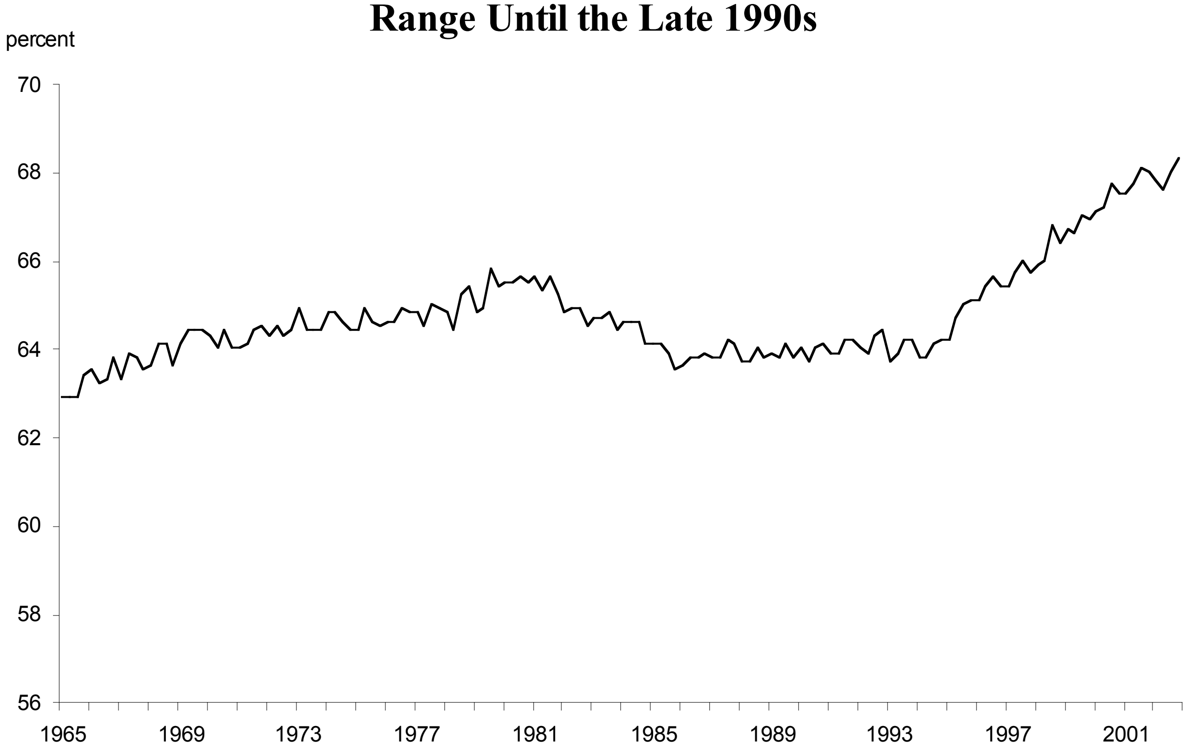

Table 1: Johansen-Juselius Estimates of Equilibrium Consumption Implied by Cointegrating Vectors
(Sample: 1955:1-02:4, uses data in levels from 1954:1-2002:4)

Models:	1	2	3	4	5	6	7	8
	Services & Nondurable Goods (c^{ns})				*Services & Nondurables + Durable Service Flow(c^{nsd})*			
	Standard	*Load*	*Standard & MEW/REFI*	*Load & MEW/REFI*	*Standard*	*Load*	*Standard & MEW/REFI*	*Load & MEW/REFI*
Non-Property Income (y)	0.6097** (13.42)	0.7608** (10.29)	0.5990** (18.80)	0.7338** (18.02)	0.5876** (10.46)	0.7387** (17.80)	0.5414** (11.74)	0.7485** (24.07)
Non-Stock-Market Wealth (w^o)	0.3235** (7.41)	0.1772* (2.14)	0.3466** (10.39)	0.1515** (2.97)	0.3625** (6.72)	0.1912** (4.17)	0.4432** (9.17)	0.1361** (3.50)
Stock Wealth (w^s)	0.0517** (8.21)	0.0212* (2.57)	0.0241** (4.90)	0.0362** (8.82)	0.0704** (8.95)	0.0381** (7.77)	0.0255** (3.53)	0.0442** (13.91)
Stock Wealth * Load (w^sL)		-0.0016* (2.38)		-0.0015** (-3.66)		-0.0024** (-6.51)		-0.0027** (-8.33)
Refinancing Savings (REFI)			14.761** (6.44)	5.7893** (2.99)			19.1693** (5.78)	4.7306** (3.21)
Mortgage Equity Withdrawal (MEW)			0.3750+ (1.78)	0.5164** (2.71)			0.7935** (2.59)	0.5674** (3.89)
Trace no CE	48.674*	74.608*	116.422**	137.468**	47.210*	82.623**	110.495**	147.088**
Trace 1 CE	24.769	46.558	67.258	93.638	23.933	46.630	62.881	93.232
Lag Length	1	1	2	3	1	3	2	3
# sign vectors	1	1	1	1	1	1	1	1

Coefficient signs from cointegrating vectors are flipped to express equilibrium consumption. Constants are omitted to save space.
t-statistics in parentheses. *(**) denotes significant at the 5% (1%) level. Critical values for trace statistics vary with the vector size.

Table 2: DOLS Estimates

(Sample: 1955:1-02:1, uses data in levels from 1954:1-2002:4)

Models:	1	2	3	4	5	6	7	8
	Services & Nondurable Goods (c^{ns})				***Services & Nondurables + Durable Service Flow(c^{nsd})***			
	Standard	***Load***	***Standard & MEW/REFI***	***Load & MEW/REFI***	***Standard***	***Load***	***Standard & MEW/REFI***	***Load & MEW/REFI***
Non-Property Income (y)	0.5544** (22.80)	0.6876** (20.88)	0.5808** (28.04)	0.6708** (29.70)	0.4990** (15.27)	0.7098** (27.81)	0.5236** (18.22)	0.7207** (39.82)
Non-Stock-Market Wealth (w^o)	0.3732** (15.23)	0.2095** (5.65)	0.3457** (15.96)	0.2244** (8.20)	0.4513** (13.64)	0.1991** (6.84)	0.4314** (15.22)	0.1768** (8.09)
Stock Wealth (w^s)	0.0524** (14.40)	0.0405** (9.15)	0.0367** (14.85)	0.0347** (13.29)	0.0652** (16.89)	0.0448** (12.16)	0.0483** (11.81)	0.0410** (17.33)
Stock Wealth * Load (w^sL)		-0.0017** (-4.90)		-0.0012** (-4.19)		-0.0026** (-8.96)		-0.0025** (-10.39)
Refinancing Savings (REFI)			11.3908** (8.47)	8.7253** (7.52)			11.1159** (6.68)	5.4791** (6.49)
Mortgage Equity Withdrawal (MEW)			0.0078 (0.07)	0.2407* (2.54)			-0.1432 (-1.12)	0.3248** (4.27)

t-statistics in parentheses. *(**) denotes significant at the 5% (1%) level.

Table 3: DOLS Estimates With Time*Stock Wealth
(Sample: 1955:1-02:1, uses data in levels from 1954:1-2002:4)

Models:	1	2	3	4	5	6	7	8
	Services & Nondurable Goods (c^{ns})				***Services & Nondurables + Durable Service Flow(c^{nsd})***			
	Standard	***Load***	***Standard & MEW/REFI***	***Load & MEW/REFI***	***Standard***	***Load***	***Standard & MEW/REFI***	***Load & MEW/REFI***
Non-Property Income (y)	0.5752** (28.19)	0.7549** (13.24)	0.5939** (33.78)	0.6850** (13.07)	0.5372** (31.86)	0.6855** (15.06)	0.5544** (35.10)	0.6561** (14.11)
Non-Stock-Market Wealth (w^o)	0.2667** (8.33)	0.2011** (5.62)	0.2710** (10.03)	0.2195** (6.74)	0.2560** (9.32)	0.2022** (6.60)	0.2548** (10.68)	0.1993** (6.86)
Stock Wealth (w^s)	0.0163^{+} (1.92)	0.0626** (5.15)	0.0155* (2.21)	0.0386** (3.17)	-0.0009 (-0.12)	0.0367** (3.53)	-0.0018 (-0.31)	0.0230* (2.20)
Stock Wealth * Load (w^sL)		-0.0027* (3.61)		-0.0014^{+} (-1.84)		-0.0022** (-3.64)		-0.0015* (-2.29)
Stock Wealth * Time (w^sT)	0.00011** (4.23)	-0.00010^{+} (-1.73)	0.00008** (3.51)	-0.00002 (-0.31)	0.00021** (9.44)	0.00004 (0.76)	0.00018** (10.22)	0.00008^{+} (1.65)
Refinancing Savings (REFI)			10.2395** (7.81)	8.5010** (5.95)			8.3943** (7.49)	6.5017** (5.61)
Mortgage Equity Withdrawal (MEW)			0.1274 (1.10)	0.2524** (2.59)			0.1395 (1.31)	0.2717** (2.84)

t-statistics in parentheses. *(**) denotes significant at the 5% (1%) level.