Has There Been a Great Risk Shift?

Trends in Economic Instability Among Working-Age Adults

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Abstract

A number of researchers—most prominently, political scientist Jacob Hacker—have argued that economic risk has shifted in recent years to workers and families from employers and government. This “great risk shift” has led to a “new economic insecurity” that demands new policy responses from government. Hacker’s primary evidence in support of his argument was a chart indicating that family income volatility had risen 200 percent from 1974 to 2002, later revised to 100 percent in response to methodological problems I discovered. These results were the latest in a long line of research on economic instability and inspired a wave of subsequent research. Most of these studies have relied on the Panel Study of Income Dynamics and have reached conclusions that differ from findings based on administrative data.

This dissertation finds that when several crucial methodological issues are addressed correctly, the Panel Study of Income Dynamics yields the same conclusions as the research using administrative data. Contrary to the risk-shift hypothesis, male earnings instability and family income instability did not rise much between the early 1980s and 2004, across a range of different instability measures. Female earnings instability fell between the late 1960s and the early 1980s and then was flat or continued declining. Other indicators of economic risk and insecurity also fail to support the risk-shift hypothesis.
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Chapter 1

Introduction

Increasingly, economic insecurity is a major concern of Americans once thought to be beyond its cold reach: middle-class professionals who have gone to college, or even beyond, but who suddenly find that their education and skills are no longer a guaranteed safety net.

By framing what usually are treated as distinct issues – pensions, health care, jobs and so forth – within a unified thesis, Mr. Hacker tells a coherent story about economic insecurity. And, by and large, the thesis is compelling.
– Roger Lowenstein, “The Economic Miracle as an Economic Mirage”

When all is said and done, the recession that began in late 2007 will almost surely be regarded as the worst since the Great Depression. In March of 2009, unemployment rose to 8.5 percent, a level not seen in over two decades. The net worth of Americans declined by 18 percent in 2008 alone. The last years of the 21st century’s first decade have understandably featured a vigorous debate about economic security, risk, and the American social contract.

Even before the financial crisis began, the “jobless recovery” wedged between the bursting of the tech-stock bubble and the housing bubble had inspired a wave of commentary arguing that the economy had become fundamentally more risky for American families. In some ways, this sense of alarm was an extension of earlier waves of insecurity inspired by stagflation in the 1970s, deindustrialization and Japanese

1 Hacker (2006d).
2 Lowenstein (2006).
competition in the 1980s, and the “white-collar recession” and downsizing of the early
1990s. But the “new economic insecurity” was typically said to be, in a word, new.3

As depicted by such writers and researchers as Jacob Hacker, Elizabeth Warren,
Peter Gosselin, and others, the United States has recently entered a period characterized
by rising economic risk, increased economic volatility, cost-shifting from employers and
government to individuals and families, and inescapable pressures to spend time at work
rather than with family, to live in the best neighborhoods with the best schools.

Among these social critics, none is more closely identified with the claim of a
fundamentally riskier economy than Hacker, and Hacker’s primary evidence for this
claim has been his own finding that income volatility has risen dramatically since the
early 1990s. This thesis examines that claim and the broader question of how economic
instability and volatility have evolved since the late 1960s. Mine is not the first attempt
to do so, but it improves upon the extensive previous literature by comprehensively
assessing a range of measures and paying meticulous attention to a number of key data
and measurement issues.

Ultimately, concern about economic instability is bound up in anxiety about
economic risk. Before delving into the details of economic instability measurement, it is
useful to first discuss the various sources of economic risk and the available ways in
which individuals insure themselves against risk. The extent and effectiveness of
insurance is one consideration in interpreting levels of and trends in economic instability.
Others include the extent to which income movements are anticipated or come as a

Dream”), Wheary et al. (2007) (“By a Thread: The New Experience of America’s Middle Class”), Warren
(2006) (“The security of middle-class life has disappeared. The new reality is millions of families whose
grip on the good life can be shaken loose in an instant.”). See also Gosselin (2008).
surprise, the importance of levels of economic well-being rather than departures from these levels, and the trade-off between insuring risk and creating moral hazard.

Sources of Economic Instability and Risk

While it is natural to think of features of labor markets as the primary force behind economic instability, the qualities that individuals carry with them upon entering the labor force may be equally or even more important. Indeed, in many respects, they are opposite sides of the same coin. If labor markets change in such a way that less-educated workers experience an increase in volatility, should we attribute the increase to market institutions or to inadequate educational attainment? People who emphasize personal responsibility and individualism will tend to fix blame on workers and their choices. Those who regard laissez-faire arguments suspiciously will point to labor markets, employers, and lax regulation. Regardless, in a society as individualistic as the United States, disadvantageous family background, genetic endowments, and pre-labor-market personal qualities are undoubtedly important sources of economic risk.

Once in the labor force, risk from labor and consumer markets is a fact of life. At the interpersonal level, risk is inherent in the dependence of most workers on their coworkers and managers or on cooperation from customers. Some labor and consumer market risk resides at the occupational or firm level, affecting workers with similar roles in the economy regardless of employer or all employees within a firm regardless of their position. Other risk may be located at the level of industry or within specific geographic areas defined by political boundaries (cities) or ecological features (areas with cold winters). And some economic risk originates in and affects the entire global economy.

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4 The following two sections draw from Heathcote, Storesletten, and Violante (2008a).
Similarly, economic risk also inheres in financial and real estate markets, as the current crisis has clearly illustrated. Investors and lenders confront risk directly in the form of delinquency and default, fluctuating asset prices, and the costs of decisions made or not made. With high home ownership and the explosion of defined contribution retirement accounts, more and more people are directly involved in capital markets. Many others are exposed to risk indirectly via the ability or inability of capital markets to promote access to credit, liquidity, and economic growth.

The possibility of health problems is another source of economic risk. Indeed, serious illness or injury may inhibit work, thereby reducing income, even as it increases medical costs. Health problems may also force workers to retire earlier than planned even as they increase the likelihood that those same workers will have inadequate resources to cover their expenses in retirement. On the other hand, the absence of health problems upon retiring presents the risk that retirees will outlive their private savings and be forced to downgrade living standards.

Finally, changes in family composition also present economic risks. Forming a household independent of parents exposes young adults to new economic risks by narrowing the extent of income pooling and removing economies of scale. Divorce (or dissolution of other romantic relationships) has similar effects. Adding children or elderly parents to a household will tend to raise expenses more than it increases income.

**Insurance against Economic Risk**

The importance of economic risk as a policy issue depends on a number of considerations. One of the most central is the question of how much insurance there is
against risk. First and foremost, individuals insure against risk in the choices they make that affect how they fare in the labor market. These choices include those related to their educational attainment and human capital investment; selection of occupation, industry, and firm; how many hours to devote to working; how much effort to put into work; and the amount of residential mobility and flexibility to accept.

Of course, what one person sees as individual “choice” looks to another like the inescapable product of opportunities or a lack thereof. Parents, for example, certainly insure against economic risks their children will face by investing time and money in their children’s development. Similarly, transfers within one’s family also provide insurance against risk, whether they take the form of financial help in rough times or the provision of free housing to post-collegiate children.

Markets also offer a number of ways to insure against economic risk. Private insurance exists to mitigate the effects of disability, death of a family member, catastrophic health problems or injuries, and loss of property, to name just a few examples. Financial and capital markets provide access to unsecured and secured credit and create opportunities for precautionary saving and wealth-building. Individual choices, including the extent of portfolio diversification and how aggressively to seek returns, interact with the opportunities that markets provide.

Other sources of insurance are matters of political economy—of the institutions, policies, and laws that societies establish to mediate when economic risks are realized. These include regulation of labor markets and private insurance and other financial and capital markets, provision of public insurance programs such as unemployment
insurance, public investments in workers and children, progressive taxation, and bankruptcy, debtor, and family law.

A final important way that people can insure themselves against economic risk is to pool resources with a spouse or romantic partner. Such an arrangement can reduce risk in two ways. First, if two or more people can achieve an adequate standard of living without all of them fully employed, then there is reserve labor that can fill in should one person run into problems. Second, when all potential workers are fully employed, in so far as their risks of economic difficulty are not perfectly correlated, then on average they will be less likely to face difficulties than if they were not pooling resources.

**Anticipated versus Unanticipated Instability**

Not all economic volatility is unanticipated. Younger adults may leave the workforce to return to school, relying on grants, loans, and savings to keep them afloat. Older adults may cease working and enter into retirement or scale back to part-time employment as they transition to retirement. Other workers may reduce their consumption and increase their savings in anticipation of leaving their job (or their spouse) or having a child. Gains in discretionary income can also be anticipated, as when a person enters or re-enters the workforce, moves in with a romantic partner, or sends an adult child out into the world.

Indeed, one can anticipate some degree of volatility far into the future in thinking about the trade-offs involved in human capital decisions. Flavio Cunha, James Heckman, and Salvador Navarro, for instance, have estimated that 60 percent of the lifetime variability in the returns to a college degree is forecastable from the information
individuals have at the time they decide whether to attend or not. Kjetil Storesletten, Chris I. Telmer, and Amir Yaron found that idiosyncratic labor income risk is strongly countercyclical, which adds empirical support to the intuition that people can observe cyclical changes in economic risk and take action to insure themselves.

Level of Economic Well-Being versus Economic Instability

Consider facing one of two situations. In either situation you will make $40,000 this year but only $30,000 next year. In the first scenario, you are partnered with someone who cannot work in either year, perhaps due to child care responsibilities, so the two of you experience a drop in income of 25 percent and end up with $30,000. In the second scenario, your partner earns $100,000 this year but only $50,000 next year. The drop in this scenario is $60,000 – six times that in the first scenario—and the percent change is nearly 45 percent instead of 25 percent. Yet you are left with $80,000. Which scenario would you prefer—less of a drop in percentage terms but lower income, or a larger drop and higher income?

The choice depends heavily on two considerations. The first relates to economists’ idea of “utility”—how much worse off would one feel without that $60,000 versus the $10,000 one would lose in the first scenario? On the one hand, the difference between $140,000 and $80,000 may be less painful than the difference between $40,000 and $30,000—perhaps because $80,000 is adequate for basic needs while $30,000 is not.

On the other hand, if people generally assume that they need not worry about an income drop, then they may make financial commitments that are difficult to extricate.

6 Storesletten, Telmer, and Yaron (2004).
themselves from. With $140,000 in income, a couple might have a correspondingly larger mortgage payment or may pay for private school for their children. While $80,000 is a substantial amount of money compared with $30,000, it may not be enough for the couple to continue their lives without considerable disruption.

This is bound up in a second consideration—the extent and liquidity of saving that people engage in. On the one hand, with $140,000 rather than $40,000, a couple might be able to increase their saving rate, since basic needs are likely to be met with room to spare. In that case, they might be able to fully insure against a $60,000 drop more easily than they would a $10,000-loss from a base of $40,000. On the other hand, if their savings rate would be similar in either scenario, then they may be equally unprepared in both.

The point is that there may be contexts in which levels of economic well being are more important than economic stability. In that case, individuals will be willing to trade off higher instability for higher incomes.

**Moral Hazard**

In a market economy, the financial fortunes of workers and families will tend to rise and fall for reasons partly—even largely—out of their control. Even sources of risk that are ostensibly matters of personal responsibility—educational attainment, say, or decisions around family formation—are often the result of cumulative youthful choices the short-sightedness of which is evident only in retrospect.

Increasingly, liberal observers have connected the idea of opportunity to security against economic risk. If people have little in the way of a safety net in the event of
failure, it is said, they will be unwilling to take risks that might improve their own
economic standing and those of others.\textsuperscript{7} But protecting citizens from market risk creates
other societal risks. In work with Christopher Jencks, I found that the federal and state
welfare reforms of the 1990s—combined with other social policy reforms that promoted
work—were largely responsible for the gains single mothers made during the economic
expansion later in the decade.\textsuperscript{8} Prior to welfare reform, Aid to Dependent Families with
Children (AFDC) had sheltered single mothers and their children against market risk
during recessions more effectively than it does today. But during expansions, it acted as
a poverty trap by discouraging many recipients from taking advantage of tight labor
markets.

In general, state-sponsored protection from market risk will often entail some
degree of moral hazard. Shielding individuals fully from the consequences of their
actions will encourage them to take action without regard to those consequences.
Sometimes that will yield benefits. For instance Daron Acemoglu and Robert Shimer
found that unemployment insurance increases unemployment but also increases
productivity by encouraging workers to seek high-productivity jobs and encouraging
employers to provide them.\textsuperscript{9} But public risk protection can also do more harm than good.
Indeed, it might be the case that poorly-designed safety nets can encourage behavior that
may be economically benign for most individuals but that might have negative effects in
the aggregate by redirecting effort and investment in inefficient ways. Alternatively,

\textsuperscript{7} Hacker and Teixeira (2006).

\textsuperscript{8} Winship and Jencks (2004a, 2004b).

\textsuperscript{9} Acemoglu and Shimer (2000).
public efforts to mitigate risk can simply not work, as would be the case if public
insurance simply crowds out private insurance.

The discussion to this point shows that some kinds of economic instability are
worse than others. Indeed, in some cases economic instability may have benefits—or
economic stability may have costs—that make it undesirable to reduce levels of risk.\footnote{See also Brown, Haltiwanger, and Lane (2006), who find that economic volatility has the effect of replacing jobs in low-productivity firms with jobs in higher-productivity ones, leading to growth in employee compensation and economic growth.}

Correctly measuring levels of and trends in economic instability does not tell us
everything we need to know about whether risk levels are too high or too low. But as
will be seen, correctly measuring economic instability is anything but straightforward in
real-world data sets.

**Measuring Economic Instability**

At a basic level, this thesis is concerned with trends in income movements.
Changes in income can produce steady or discrete movement in one direction, up or
down, or it can involve a path that bounces around in both directions. These distinct
concepts can be thought of as “mobility” and “volatility”, respectively. A worker whose
income bounced around randomly from year to year would experience significant
volatility, but comparing two years might not reveal much mobility. On the other hand,
most workers might be progressing upward or downward along steady trajectories, in
which case there would be little volatility but significant mobility.\footnote{Nichols (2008) and Nichols and Zimmerman (2008) make a similar distinction.}

Nearly all measures of income change examined to date can be grouped into those
looking at (1) how often people experience sudden sizable shifts in where they stand
relative to others, (2) how often workers experience sudden large income gains or losses, (3) the extent to which income differences at one point in time are associated with differences at a later point in time, (4) the full distribution of sudden gains and losses, (5) the distribution of the typical individual’s income within a short window of time, and (6) the range of income shocks across individuals.

The first three of these types of measures have been most common in mobility research. There is a vast and long-standing research literature on economic mobility, most of it focused on intergenerational mobility, but some of it examining mobility within individuals’ lifetimes. The research on intra-generational mobility includes studies looking at mobility over periods of ten years or more as well as work focused on shorter time spans. The wider the time interval within which mobility is examined, the more mobility studies reflect career and aging dynamics rather than short-term income fluctuations.

This thesis examines mobility over time spans of no more than five years, which I refer to as “short-term mobility”. Research on short-term mobility can be divided into studies examining relative mobility (movements in the position, or rank, occupied in the income distribution relative to others), those examining absolute mobility (income movements up or down without reference to relative positions), and those examining mobility as time independence (lack of association of incomes over time).12

12 This classification differs somewhat from that of Fields and colleagues (who distinguish between two types of relative mobility and two types of absolute mobility) and those of researchers who distinguish only between absolute and relative mobility (generally treating Fields’s time-independence as relative mobility). For Fields’s classification, see Fields and Ok (1999), Fields, Leary, and Ok (2000), and Fields (2004). For a more general discussion of relative and absolute mobility, see Sawhill and Morton (2007). I thank Fields for providing me with copies of his papers.
Measures of absolute mobility do not consider individuals’ positions in the earnings distribution; they consider any decline or increase in earnings as a change in mobility, regardless of how others are doing. Relative mobility can remain stable while absolute mobility increases or decreases if everyone experiences the same change in earnings.

Measures of mobility summarize either directional or nondirectional movement. That is, they can summarize the amount of either downward or upward mobility, or they can summarize the amount of mobility regardless of direction. A primary strength of directional measures of relative and absolute mobility lies precisely in their ability to distinguish between upward and downward movements. Measures of nondirectional mobility and most measures of volatility do not convey clear information about whether income declines outnumber or dominate income increases. This shortcoming makes it difficult to interpret trends in such measures, since increases may reflect a general rise in living standards or growing inequality rather than increased risk.

_A priori_, one might expect directional relative and absolute mobility trends to follow similar but not identical paths. During recessionary periods, a relatively large number of people will tend to experience downward movement in income measured in absolute terms, and upward movement will be more common during expansionary periods. However, relative mobility should tend to be less cyclical, since ranks need not change much when many people are moving up or down. For example, the fraction of the population changing income quintiles between two years should be less sensitive to business cycles than the fraction with a change in income of 25 percent or more.
Nondirectional measures of relative and absolute mobility may show even less cyclicality than directional measures of relative mobility. That would be the case if total mobility—upward or downward—remained roughly constant over the business cycle according to some measure, with increases in downward mobility during recessions completely balanced out by increases in upward mobility during expansions.

The same would be true for measures of intertemporal income association, which are also nondirectional mobility measures. While many researchers classify all of these measures of association, such as regression coefficients using logged earnings and autocorrelations, as indicators of relative mobility, this is not strictly true. A Pearson correlation coefficient, for instance, can be less than 1.0 even if no one’s ranking changes, and it can change over time without ranks changing. Absolute mobility is necessary for the correlation coefficient to indicate mobility, but it is not sufficient. Relative mobility is sufficient for the correlation coefficient to indicate mobility, but it is not necessary.\(^{13}\)

Absolute mobility measures that examine the likelihood of experiencing a change in income of a given size look at only part of the distribution of income changes. A fourth line of research on economic instability involves summarizing the dispersion of the entire distribution of income changes. Greater dispersion of changes, in this approach, constitutes greater mobility, since it implies that relatively large income changes are becoming more common and relatively small income changes less common. If one measures absolute mobility as the mean across individuals of the squared change in income between two years, there is a clear link to the variance of income changes, which

\(^{13}\) On the other hand the Spearman rank correlation coefficient or the centile correlation coefficient cannot change without rank mobility changing, so they are properly viewed as measures of relative mobility.
is the mean across individuals of the squared change in income after centering (or de-
meaning) incomes.

A downside of this approach is that the dispersion of income changes can grow
simply if income gains become bigger or more common or if inequality is increasing.
Not only that, but it fails to differentiate between a scenario in which the typical person is
facing bigger income changes over time and one in which the typical income change
stays the same over time but changes increase for those who experience the largest
changes.

The fifth type of income-change measure most closely corresponds with the idea
of volatility. It measures within-person income dispersion over several years. Such
measures have the benefit of showing greater volatility when individuals’ incomes
fluctuate over several years, rather than when individuals simply experience a single
large decline or increase. On the other hand, within-person income dispersion will also
increase if incomes are moving steadily in one direction or another rather than bouncing
around. If one measures within-person dispersion as the variance of a person’s income
over two years, the mean dispersion across individuals equals one-fourth the mean
squared income change across individuals. Using a larger number of years to measure
within-person dispersion adds information about income volatility, since there is likely to
be more fluctuation across several years than across two.

The final type of measure comes from a distinct line of research modeling
earnings dynamics formally and produces parametric estimates of the dispersion of
income shocks. In these studies, earnings are separated into permanent and transitory
components, sometimes defined conditional on observable demographic characteristics.
Shocks to earnings operate either through the transitory component (in which case they are generally modeled as white noise, sometimes persisting to affect subsequent years) or, less often, the permanent component (in which case they are modeled as depending on past permanent shocks). Volatility is equated with the dispersion of these shocks. Typically, a model is specified that implies restrictions on earnings variances and within-person covariances (or the variances and covariances of earnings changes). Minimum-distance methods are then used to choose the model parameter values that produce an earnings covariance matrix that most closely resembles the actual covariance matrix.

By distinguishing permanent and transitory components, these studies do a better job analytically of addressing the difference between mobility and volatility than other measures. But of course, all models are simplifications, and if they do not describe reality well enough, then the estimates of volatility they produce may be biased.

In fact, the simplest models used effectively rule out enduring mobility, modeling earnings as consisting of a never-changing permanent component and a sequence of discrete random shocks, the effects of which disappear with the next shock. Most models impose the same baseline income profile on all individuals within particular demographic groups, ruling out any initial income variation based on ability, preferences for leisure, or time horizons that are not absorbed in observable covariates. Only a few allow parameters to interact with age, so that the level and dispersion of shocks can change over the life course. None allow the distribution of shocks to vary with the business cycle or include aggregate shocks corresponding with cyclical macroeconomic patterns or industry-specific trends.
Developing models of earnings and income dynamics is an important and necessary endeavor for answering many micro- and macroeconomic questions. Ultimately, they are probably indispensable for interpreting what to make of trends in income movements. However, for purposes of describing such trends, estimates based on these models abstract from the income changes that real individuals experience. Relying on them for description is akin to describing trends in income using income predicted from a regression model rather than self-reported income.

In the chapters to follow, I compute trend estimates using each of these six conceptualizations of economic instability, but I focus on two measures in particular. First, I emphasize the risk of experiencing an absolute drop in income (of 25 percent), or downward absolute mobility. The research on economic instability has tended to obscure the fact that what constitutes a risk to individuals is not volatility per se, but downward income movements. Indeed, for a given level of downward mobility, volatility is beneficial—it implies recovery from income drops. I focus on absolute mobility rather than relative mobility because what is disruptive to people in the short-run is not how one’s rank changes, but how one’s level of resources changes. Over longer periods, or across generations, there may be contexts in which relative mobility is as important as absolute mobility or more so.

The second measure I emphasize is a new indicator I call “pivot volatility”. This measure is similar to conventional measures of within-person dispersion in beginning with income movements within a short-term window experienced by individuals. The experience of the typical person is then summarized. Where it differs from measures of within-person dispersion is that it explicitly determines the frequency and magnitude of
income reversals within a person’s window. Volatility occurs when reversals involve large movements (up and then down, or vice versa) and when there are multiple reversals. Pivot volatility addresses the shortcoming shared by the volatility measures used to date: their failure to distinguish between income growth and inequality on the one hand and true volatility or instability on the other.

Organization of the Thesis

In Chapter Two I examine trends in individual earnings instability. There is a vast literature related to this topic, some of it in the tradition of mobility research, some conducted by labor economists primarily interested in modeling earnings dynamics, and some of it pursued in the course of developing macroeconomic models of the U.S. economy. I bring these literatures together for the first time and attempt to synthesize the existing research, considering trends for men and women separately. I then produce my own estimates that try to improve on earlier research through more careful methodological decisions and summarize how trends using different measures compare with each other.

Chapter Three pursues an analogous aim, this time focusing on trends in household income instability. This literature is smaller and dominated by a few recent studies that have gained prominent attention. I compare my estimates to previous research on earnings and income instability and to analogous trends in earnings instability.
In Chapter Four I restrict the analyses to downward absolute mobility and examine trends by demographic group. Specifically, I consider how levels and trends differ depending on family composition changes, age, educational attainment, and race.

Finally, the concluding chapter presents trends in other indicators of economic risk and insecurity, discusses the politics of economic instability and insecurity, and provides policy options for addressing economic risk. I propose a program of “citizen benefits” to supplement and replace the current system that delivers benefits via employers. The proposal is voluntary, relatively low-cost, and designed to address a number of problems related to economic instability in a politically effective manner. It is offered as an alternative to more ambitious agendas that in my view mis-read the evidence on both economic instability trends and public opinion related to economic insecurity.

Above all, my citizen benefits agenda is offered in the belief that one can believe a problem merits public attention and a commitment of public resources without having to believe it is getting worse, and that national problems can be addressed in ways that take serious the complexity of public opinion and the virtues of markets. Above all, my analyses and my proposal try to respect the most important principle of social scientific practice: that facts matter for successfully remedying social problems.
Chapter 2

It Ain’t Got That Swing: Trends in Earnings Instability and Volatility

In the typical non-elderly American household, nearly all income comes from earnings, the vast majority of it from wages and salaries.\footnote{In the median household with a head age 20-59 in 2006, earnings accounted for over 95 percent of the household’s income, including self-employment earnings. Just 13 percent of households had any self-employment earnings. Author’s computations using the 2007 Annual Social and Economic Supplement to the Current Population Survey.} Understanding whether or not economic instability has increased, then, begins by examining trends in individual earnings mobility and volatility. This chapter provides new trend estimates using the measures typical of each line of past research. The results are all based on a common dataset and common methodological choices in order to determine where consistent conclusions can be drawn. I provide trends for men and women separately and combined. For several measures of mobility and volatility, my estimates for women are the first of their kind, and I provide the longest time series to date for many other measures. I also provide a new measure of “pivot volatility” that captures the concept of volatility better than any measures used to date.

My results are broadly consistent across measures, and with the bulk of previous research, though they add important context that helps explain differences across the research. I find evidence that the risk of a large earnings drop increased very modestly over the 1970s and early 1980s among men, but that there has been no secular increase since then. While volatility may have been greater for the self employed, earnings volatility probably increased by 20 to 30 percent among wage and salary workers, which
translates into minimal change in the experience of the typical worker. All of this change occurred before the mid-1980s. Earnings instability has declined among women. There is little indication of a “risk shift” since the early 1980s.

**Previous Research**

The research on earnings instability spans a number of distinct research literatures, including mobility scholarship, studies of economic insecurity, efforts to model earnings dynamics, and macroeconomic modeling using heterogeneous agents. Summarizing the literature and making sense of it—identifying and explaining inconsistencies—is a formidable challenge. In Appendix One, I present an extensive review of previous research on earnings instability. I summarize my conclusions from that review here.

*Short-Term Directional Mobility*

A number of studies examine trends in short-term downward and upward mobility, measured in either relative or absolute terms. Short-term relative mobility involves upward or downward movement in rank within the earnings distribution over a period of five years or less. For most measures used, upward and downward relative mobility need not balance each other out, since they are typically defined in terms of moving from a top or bottom quantile into any of several other quantiles. Mobility out of the bottom quintile, for instance, need not be matched by mobility from the top quintile, because there are movements from the middle three quintiles as well.
Short-term relative mobility declined during the postwar period, through the early 1960s, with downward mobility declining through the mid-1960s (Kopczuk et al., 2009). Upward and downward mobility followed cyclical patterns from the 1970s forward, with upward mobility trending downward from the mid-1970s to the early 2000s but downward mobility changing little beyond a drop in the early 1980s. Kopczuk et al.’s results did not disaggregate men and women. Buchinsky and Hunt (1999) found declines in downward relative mobility from the top of the earnings distribution and upward relative mobility from the bottom over the 1980s for men and women.

Turning to absolute mobility—the likelihood of experiencing a large earnings gain or drop over five years or less—Dahl et al. (2007) find that downward mobility also declined during the first half of the 1960s and then increased, declining over the decade as a whole. Downward absolute mobility probably increased among men in the 1970s (Dahl et al., 2007; Dynan et al., 2007; Hacker, 2007). However, the 1980s featured little change in upward or downward mobility, or perhaps declines in one or both. Results from the 1990s are inconsistent, but it appears that there was either little change or a decline in downward absolute mobility at the same time that there was little change or an increase in upward absolute mobility. The first years of the current decade, marked by the 2001 recession, featured increases in downward absolute mobility and declines in upward absolute mobility. As with directional measures of relative mobility, downward and upward absolute mobility show strong countercyclical patterns. The results of Dahl et al., using Social Security administrative data, conflict with the PSID-based studies of Dynan et al. and Hacker over the period as a whole, with Dahl et al. finding a decline in
downward and upward absolute mobility from 1980 to the early 2000s instead of an increase.

**Short-Term Non-Directional Mobility**

As with directional relative mobility, short-term non-directional relative mobility declined from the late 1940s through the early 1960s, then began increasing, probably reversing the early ‘60s decline, though there are inconsistent findings at the end of the 1960s. All but one study find increases in non-directional mobility over the 1970s. Non-directional mobility declined during the 1980s and in the early 1990s before flattening out through the early 2000s.

Non-directional absolute mobility increased in the 1970s among men, but most studies show flat or declining mobility over the 1980s. Trends thereafter are inconsistent. Studies using measures of intertemporal earnings association—correlations between incomes measured one or a few years apart—are remarkably inconsistent, and little can be said about how the correlation has changed since the late 1960s. Prior to 1960, Kopczuk et al. (2009) found very large declines in mobility, but we have no other pre-1960 studies, so we cannot be sure their results would hold up using other data or methods.

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3 Ibid. Fields, Leary, and Ok (2000); Daly and Duncan (1997). Moffitt and Gottschalk (1995) is the exception.


5 Dahl et al. (2007); Dynan et al. (2007); Jensen and Shore (2008); Fields, Leary, and Ok (2000).
Dispersion of Earnings Changes

Rather than focus on trends in the likelihood of experiencing a change of a given size, up or down, in earnings, other studies look at the entire distribution of earnings changes and see whether it has widened over time. Most research finds that male earnings instability measured this way increased during the 1970s.\(^6\) Dynan et al. (2008) find a small decline for women.

The research also tends to find increases in male earnings volatility during the early 1980s recession, followed by declines over the mid-1980s.\(^7\) It disagrees as to whether volatility rose or fell over the decade as a whole. According to Dynan et al.’s and Dahl et al.’s results, earnings movement among women declined steadily during the 1980s.

Male earnings volatility rose temporarily during the early 1990s recession, but there is little consistency across studies in the trends found over the rest of the decade, except that it rose late in the decade and in the early 2000s.\(^8\) Dahl et al. and Dynan et al. found declines among women over the 1990s and early 2000s. As was the case for the research on short-term directional mobility, the Dahl et al. study disagrees with the PSID-based studies after the early 1980s in finding little change in male volatility. The research generally agrees that this measure of volatility follows a cyclical pattern similar to that found for directional mobility.


\(^7\) Moffitt and Gottschalk (2008); Dynan et al. (2008); Shin and Solon (2009); Dynarski and Gruber (1997); and Cameron and Tracy (1998); Dahl et al. (2007).

\(^8\) Shin and Solon (2009); Dahl et al. (2007); Moffitt and Gottschalk (2008); Dynan et al. (2008).
**Within-Person Earnings Dispersion**

Measuring how the typical person’s earnings vary over time corresponds more closely with the concept of volatility than the above measures do. Among men, within-person dispersion of earnings grew over the 1970s and the 1980s, with counter-cyclical increases in the early 1980s and the early 1990s. There was probably little change thereafter, though the various studies are not entirely consistent.9 Among women, Gosselin (2008) found that volatility was flat in the 1970s and declined in the subsequent decades. Studies combining men and women consistently show increases in volatility in the 1970s, but there is little agreement about what happened thereafter.

**Across-Person Dispersion of Earnings Shocks**

As noted in the introduction, a sizable literature models earnings dynamics as processes subject to “shocks”, and dispersion in these shocks constitutes another measure of volatility. Nearly all of the research in this tradition focuses on male heads, and nearly all of it relies on the PSID. Volatility by these types of measure was either flat or increasing over the 1970s. Most studies find rising volatility in the 1980s, though several find a flat or declining trend. Most notably, two studies that use Social Security Administration data rather than the PSID show a decline over all or part of the 1980s (Gottschalk and Moffitt, 2007; Mazumder, 2001). The research generally agrees that volatility rose in the early 1980s recession and declined in the mid-1980s. The research also consistently shows that volatility increased during the recession of the early 1990s and over the course of the decade. The evidence is inconsistent as to whether it increased

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9 Moffitt and Gottschalk (2008); Gottschalk and Moffitt (2007); Daly and Duncan (1997); Comin et al. (forthcoming); Gosselin (2008).
in the early 2000s. Among the studies that span the years in the PSID, male earnings volatility increased by 15 to 65 percent from the early 1970s to the early 2000s. Kopczuk et al.’s (2009) results imply that volatility was higher before 1960 than it was in subsequent decades.

Summary of Previous Literature

The 1970s appear to have been a bad decade for men: more downward earnings mobility and more volatility. Volatility may have also increased among women. Since 1980, downward and upward mobility have followed a cyclical pattern, and the secular change has been minimal. However, volatility as measured by the within-person dispersion of earnings or by dispersion of shocks to earnings appears to have increased between 1980 and the early 2000s.

This conclusion, however, relies on evidence that may be incomplete in several ways. First, the PSID-based studies, which constitute much of the research, rely on a survey with a number of features that greatly complicate its use in research on volatility and mobility, and which has often produced findings that differ from those obtained from Social Security Administration data. As a result, PSID-based results may not accurately portray trends in earnings instability. This could be true either because of problems with the PSID data itself or because of the methodological decisions that past researchers have made. My analyses in this chapter attempt to improve on the methodological decisions used by previous researchers and to standardize them across different types of mobility and volatility measures.
The second question is whether the within-person dispersion measures typically used or the model-based measures of dispersion of shocks adequately differentiate between trends in earnings growth and trends in volatility per se. Finally, the existing research is sparse in certain aspects. Most obviously, there is much less research on trends among women than among men. Furthermore, for some measures there exist no time series that include estimates for men and women separately as well as combined. In addition, some of the literatures do not extend as far back as the data allows or as far forward. The estimates I present here are an attempt to fill in these gaps.

Methods and Data

My goal for this chapter is to see what internally and externally consistent conclusions may be drawn about trends in earnings movements when various measures used in previous research are examined using a single dataset, both aggregating and disaggregating men and women, applying consistent methodological decisions, and taking great care in making those decisions.10 All of my STATA programming files and my PSID extract are publicly available at www.scottwinship.com. To anticipate a major result of my experimentation, it is particularly important for consistency of trends that very low earnings be appropriately trimmed. Furthermore, there appear to be sound theoretical reasons for excluding individuals with no reported earnings from analyses for purposes of understanding trends in volatility and instability.

10 The numerous tests that I conducted to examine the sensitivity of my estimates to methodological decisions, noted throughout the discussion to follow, were conducted in the course of producing volatility estimates based on the variance decomposition model of Gottschalk and Moffitt.
Data

Like most of the studies reviewed in Appendix One, I use the Panel Study of Income Dynamics (PSID), conducted by the University of Michigan’s Institute for Social Research. The PSID interviewed a representative sample of nearly 3,000 households from the lower 48 states in 1968 (the “SRC sample”), along with another sample of nearly 2,000 low-income non-elderly households from metropolitan areas and southern rural areas (the “SEO sample”). Until 1997, surveyors conducted annual follow-up interviews with these respondents, tracking families as they moved, incorporating new family members as needed, and following adult children who left home to form their own households. Over the course of 1997 and 1999, PSID administrators added a sample of 511 families that had immigrated to the U.S. after 1968 in order to make the survey representative of the 1997-99 population. The PSID also went biennial after 1997, with surveys in 1999, 2001, 2003, and 2005. Because analysis of PSID data must overcome a number of subtle and obscure pitfalls, I discuss the data in detail in this section to facilitate greater consistency in future research on earnings and income changes.

PSID users must first decide whether to confine themselves to the SRC sample or use the entire “core” sample, which includes the SEO and immigrant samples. The PSID data includes weights that are designed to adjust for differential sampling probabilities and attrition in the entire core sample. Using the weighted core sample has both benefits and costs. On the benefit side, it maximizes sample size and adjusts for

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12 The PSID also includes a “Latino sample”, which was surveyed from 1990 until 1995, when it was dropped. The core sample does not include the Latino sample.
attrition. It also adjusts for other changes in the PSID sample from year to year. For instance, in 1990, 1992, 1993, and 1994, surveyors made concerted recontact efforts that resulted in the successful re-incorporation of several thousand attriters back into the survey. In 1990, 1993, 1994, 1996, 1997, and 2005, the PSID’s rules for following household members that move were adjusted, and in 1994 the sample was redefined to include additional children. The SEO sample was also reduced for budgetary reasons in 1997. These changes produce “seams” in the data that make year-to-year comparisons potentially problematic.\(^{13}\)

Using the entire core sample also has costs. The initial selection of the SEO sample involved significant departures from random probability sampling.\(^{14}\) Even absent this problem, using weights introduces inconsistencies into the data. The weights are designed with some correction for attrition, with adjustments made every five years from 1969 to 1989, in 1999, and in 2005.\(^{15}\) But because people enter and exit the PSID sample from year to year, using weights designed for year \(t\) to estimate earnings changes between years \(t\) and \(t'\) will not adjust the sample correctly.

Unequally spaced recontact efforts and introduction of a post-1968 immigrant sample make this issue even more problematic. From 1993 to 2003, respondents who were successfully added as part of the 1993 and 1994 recontact efforts were given weights only if they had been present in the 1989 wave. But in the 2005 data, weights


\(^{14}\) Brown (1996).

\(^{15}\) Heeringa and Liu (2001) and Gouskova et al. (2007). Gouskova et al. indicate that an attrition adjustment was made in the yet-to-be-released revised 2003 weights, but the 2005 weights are the only publicly available ones to reflect this adjustment. In the revised weights, for 1993 to 2003, there is no 1999 attrition adjustment.
were assigned to these respondents (the PSID weights were recently adjusted so that the 1993 to 2003 weights are consistent with the 2005 ones, but few papers cited in Appendix One were written before this change was made).\(^{16}\) Furthermore, with the incorporation of the immigrant sample, the weights use post-stratification to make the PSID sample nationally representative. Prior to 1997, the sample in any year is supposed to be representative of noninstitutionalized Americans alive in 1968, plus their descendants. From 1997 onward, the weighted PSID sample is supposed to be representative of the noninstitutionalized U.S. population for a given year.\(^{17}\)

To determine which option would provide the most accurate results, I compared trends in total earnings variances in the Annual Social and Economic Supplement to the Current Population Survey to trends using the SRC sample and the full weighted core sample.\(^{18}\) Looking at levels and trends in the variance of male wages, male earnings, and female earnings, the SRC estimates were consistently as close or closer than the weighted core results to the CPS results (see Figure 2.1 for the SRC vs. CPS comparisons). For this reason, I chose to use only the SRC sample in my analyses. That said, I also ran my variance decomposition models using the weighted core sample, and the differences were not large.

\(^{16}\) Gouskova et al. (2007).

\(^{17}\) Heeringa and Connor (1999).

\(^{18}\) I obtained the CPS data from Unicon Research Corporation (www.unicon.com). I compared the PSID variances for male family unit heads’ wages, male labor income (family unit heads and spouses), and female labor income (family unit heads and spouses) with CPS variances for male heads’ and spouses’ wages and earnings, and female heads’ and spouses’ earnings. I looked at variances in the CPS defining headship both at the household and family levels. I also used different trims of the top and bottom of these earnings distributions. Figure 2.1 presents results using 2% top and bottom trims. Full results available from the author upon request.
The PSID asks detailed questions on sources and amounts of income received by a “family unit” head and the head’s spouse or partner, if one is present. The PSID nearly always designates the male partner of a couple as the head, apparently deviating from this convention only if he is incapacitated. The earnings and income questions are fairly consistent over time, but several changes deserve mention.

Figure 2.1. Total Earnings Variance, CPS vs. PSID

Source: Author’s computations using PSID data and data from the March CPS (purchased from Unicon). CPS data is for household heads and their spouses, PSID data is for family unit heads and spouses.

19 In the PSID a family unit is defined as “a group of people living together as a family. They are generally related by blood, marriage, or adoption, but unrelated persons can be part of a FU if they are permanently living together and share both income and expenses.” (http://psidonline.isr.umich.edu/Guide/FAQ.aspx#90) There are a small number of households with more than one family unit. I ran my variance decomposition models using only family units where the head had the highest earnings among the multiple family units in a household, and the results were barely changed.

Most earnings and income variables in the PSID are aggregates of component variables. Until 1976, a number of these variables are available only in categorical form, and this is true of most of them prior to 1970. From time to time the composition of aggregate measures changed, generally as component variables were split into two or more variables. This affects the measures of heads’ and “wives” labor income. From 1968 to 1983, income from self-employment appears to have been allocated between labor and asset income using a consistent approach. The approach also appears consistent from 1984 through 1992, with some labor income potentially allocated to “wives” for the first time. In 1993, however, the approach changed. And in 1994, not only did the approach change again—with the portion of market gardening income that was allocated to asset income in the past now going toward labor income, and with the portion of roomer/boarder income that was allocated to labor income in the past now going toward asset income—but labor income from businesses and farms was removed entirely from the PSID “labor income” variables for heads and wives.

I added the labor part of business and farm income back into labor income for the post 1993 surveys.²¹ In addition, I estimated results that added the asset part of market gardening income back in and subtracted out labor income from roomers for all years. Finally, because of the relatively arbitrary allocation of self-employment income between asset and labor income, I also estimated results assigning all self-employment income to

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²¹ I had to create heads’ and wives’ labor income from business for 1997 using the edited non-Income-Plus data. For details, see the “fam93.txt” file at ftp://ftp.isr.umich.edu/pub/src/psid/documentation/93famdoc.zip.
labor—including the asset part of heads’ and wives’ farm and business income. Again, the results were largely the same.\textsuperscript{22}

The top-coding (and sometimes bottom-coding) of earnings and income component variables also changed from time to time, as did the top-coding and bottom-coding of the aggregate variables themselves. I addressed these inconsistencies by trimming the top and bottom two percent of positive earnings within age/year cells for my measures (within age/year/sex cells for estimates combining men and women).\textsuperscript{23} It would be preferable to trim the component variables prior to aggregating them, but recreating the component variables is often difficult, because they sometimes depend on multiple variables themselves, and one loses imputed values even where it is straightforward to re-create components. Any effect of changing top- or bottom-codes in these component variables should show up by comparing results for earnings measures made up of several components (such as male labor income) to results for measures with few components (such as male heads’ wages).

Another source of inconsistency is the 1993 shift to more detailed questions for certain income component variables. Rather than the initial two-question format asking if income of a certain type was received and if so, what the amount was, a more elaborate sequence was introduced. If a respondent indicated receiving income of a certain type, he or she was asked about an amount, then whether that amount was received annually, monthly, weekly, or on some other schedule, then in which specific months the

\textsuperscript{22} Shin and Solon (2009) note that the definition of heads’ wages changed in 1993, with “income from extra jobs” separated out from that year forward. But when they added this component back into wages, it barely affected the volatility estimates at all. I also found similar results when I added the component back in.

\textsuperscript{23} I relied on my comparisons with CPS earnings variances to determine what trims to use. The age categories used for the trimming were 21-30 years, 31-40 years, 41-50 years, and 51-60 years. The influence of outlying high earnings is also minimized by using the natural log transformation, which makes my results relatively insensitive to large earnings and large changes in earnings.
respondent received the income. This new sequence requires imputation of an annual amount if the respondent indicates any periodicity other than annually or monthly.

Imputation procedures also changed in the 1990s, as data collection, processing, and editing procedures were updated. Computer-assisted telephone interviewing (CATI) was introduced in 1993. Changes in the software used to collect, process, and edit the income data create potential inconsistencies in the data in 1994, 1995, 1999, and 2003, with the biggest potential seams between 1992 and 1993, 1994 and 1995, and 2001 and 2003 (which divide the period according to the version of data collection software used).^24

The new income processing software led the Institute for Social Research to release supplemental “Income Plus” files for the years 1994 to 2001 that contained updated income variables that had undergone enhanced editing. After 2001, the income processing software had improved enough that supplemental enhanced files were deemed unnecessary.^25 I use the Income Plus versions of the variables in my analyses. As a check against the potential problem of imputations of missing values, I conducted the

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^26 The Income Plus file in 1997 excluded the new immigrant sample, but edited income variables for immigrants are included in the data for that year available on the website. In the variance decomposition models of transitory earnings variance I ran using the full weighted core sample, I used these edited variables for immigrants (heads’ wages, heads’ labor income, wives’ labor income, farm income, heads’ and wives’ taxable income, others’ taxable income, money income, heads’ and wives’ transfer income, and others’ transfer income).
variance decomposition analyses excluding incomes that incorporate imputed values. The results were not materially affected.\textsuperscript{27}

Because so many methodological and data changes were made in the PSID in the early 1990s, volatility trends from the PSID that differ from other surveys in this period should be viewed as probable methodological errors. Research by the Institute for Social Research has shown that the introduction of CATI and new income processing software in 1993 increased income variances (with the income processing software probably more important than the switch to CATI). Compared with the CPS, income variance in the 1993 survey is suspiciously large (and greater than in adjacent years in the PSID). Percentiles below the median show one-time declines, and those above the mean show increases.\textsuperscript{28} From 1992 to 1996 PSID income estimates at the first, third, and fifth percentiles are below the estimates from the CPS, though the opposite is the case for all other years between 1967 and 2004 and for all but the lowest percentiles even between 1992 and 1996.\textsuperscript{29} Dynan et al. (2008) discovered a sizable jump beginning in the early 1990s in the frequency of heads reporting $0 in earnings at the same time they report working more than 120 hours. Nichols and Zimmerman (2008) also report a similar jump in the early 1990s.

Kim and Stafford (2000) enumerate the data issues during the 1990s that they urge users to take into consideration. They note that the CATI instrument used from 1995 to 2001 improved the accuracy of the data relative to the previous CATI instrument

\textsuperscript{27} Shin and Solon (2009) report that the Institute for Social Research indicated to them that the relatively large number of $1 wages in the data from 1994 onward should be treated as missing values requiring imputation. My trimming of the data apparently obviates this problem; when I estimated results dropping all $1 reported wages, the trends barely budged.

\textsuperscript{28} Kim and Stafford (2000), Gouskova and Schoeni (2007).

\textsuperscript{29} Gouskova and Schoeni (2007).
and the pencil-and-paper surveys that preceded it. They also indicate (in 2000) that the CATI instrument for 2003 onward “will be better” than the 1995-2001 instrument. The review concludes with the warning, “In brief summary, the PSID has undergone a great number of changes 1992 – 2000 and will be going through many more significant changes in the next few years. During this change process a large number of potential data seams could have arisen.”

Beyond any issues with data comparability since the early 1990s, the Institute for Social Research reports that PSID income shows declines at percentiles below the median in the late 1980s that are not present in the CPS, as well as flatter increases above the median. At the very bottom of the income distribution (below the tenth percentile), the PSID and CPS trends do not match up very well between the mid-1980s and the early 1990s, with the PSID percentiles declining anomalously. There is also an up-tick in income at these low percentiles in 2004 that is not present in the CPS.³⁰

Sample attrition and measurement error are also concerns in the PSID. Past research has found little attrition bias in the PSID, even though attrition rates are quite high.³¹ There has been less examination of measurement error. Changes in the survey instrument and income processing software were meant to reduce measurement error, but the length of the survey interview doubled between 1995 and 1999, which would be expected to increase measurement error. These possibilities are important because measurement error can affect volatility estimates. Classical measurement error inflates

³⁰ Ibid.

³¹ Fitzgerald, Gottschalk, and Moffitt (1998a); Fitzgerald, Gottschalk, and Moffitt (1998b); Lillard and Panis (1998); Zabel (1998); Beckett et al. (1988). See Nichols and Zimmerman (2008), however, for evidence that from year to year, attriters are different from those included in volatility samples in terms of the joint distribution of a number of demographic variables.
measured volatility levels, so if measurement error in the PSID declined over time, this
trend will understate the increase in volatility or overstate the decline. The opposite
would be true if classical measurement error has increased over time. If the measurement
error in the PSID is non-random, however, the effect on volatility trends is indeterminate.

Additional Methodological Considerations

In most analyses I transform earnings by taking their natural log. Using natural
logs has the advantage of transforming earnings to a scale that is mean-independent. An
across-the-board 10 percent change in logged earnings will not produce a change in the
logged earnings variance. An enormously important problem with the log
transformation, however, is that it has the effect of increasing the influence of changes in
very low earnings, while reducing the influence of changes in large earnings. A non-
negligible number of PSID respondents report annual earnings under $500. Year-to-year
changes in measured volatility can be strongly influenced by changes in the proportion of
individuals reporting such low earnings. In the PSID, there is enough year-to-year
variation in low reported incomes that the problem can drive estimated changes in (log)
volatility over time.32

Trimming earnings sufficiently before estimating volatility addresses this
problem. While trimming may remove individuals with some of the biggest earnings
increases or declines from the sample, there is no reason to think that the trend in
earnings instability is biased by applying the same trim across all years. Essentially my

32 In my early research on household income volatility, discussed in detail in the next chapter, I discovered
that this problem largely explained the extraordinary growth in income volatility that Jacob Hacker (2006a)
reported in the early 1990s (c.f., Hacker, 2008). Winship (2007) and personal communication with Hacker
(2007).
results describe instability trends for people whose earnings always put them in the middle 96 percent of earners in the years that go into the computation of a particular instability measure.

My analyses include only persons with positive earnings in the years considered. This restriction is partly for methodological reasons – the log of a non-positive number is undefined – but also because persons who go an entire year without any earnings are likely to have done so for reasons that make their inclusion in these analyses inappropriate. Setting aside people who simply misreport their earnings, they are likely to be sick or disabled, independently wealthy, homemakers, new parents, retirees, or students. Most concern over volatility trends revolves around what such trends imply for the structure and strength of labor markets.

Nevertheless, some persons without any earnings are discouraged workers, and volatility analyses should ideally retain such individuals. However, there is little reason to think that they represent a large fraction of those who go an entire year without earnings. One can gain some purchase on this question by looking at persons out of the labor force in the March CPS for 1993, the last year in which detailed questions are available. I focused on household heads and spouses between the ages of 25 and 59 who had not worked in over a year. Among women, two-thirds said that they were keeping house and were not currently interested in a job. Another 18 percent said they were not interested in a job for other reasons; half of them said they simply had no desire to work.

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33 Including persons without earnings in the variance decomposition models is problematic, even if their earnings are recoded to $1 or some other value. The model requires permanent and transitory components to sum to total earnings. A person with $0 in earnings, must have both $0 in permanent earnings and $0 in transitory earnings or must have a transitory component that exactly cancels out his or her permanent component. The distribution of transitory earnings, in expectation, should be the same among those who do not work as it is among those who do. When one attempts to estimate the variance decomposition model by including $0 earners (after recoding their earnings to be positive), the transitory variance estimates balloon. Error components models are similarly sensitive to recoding of $0 values.
while most of the rest were sick or disabled. Of the remaining 15 percent, a third were ambivalent about whether or not they wanted a job (most of them homemakers). Only 9 percent said they wanted a job, and just 2 percent both wanted a job and cited problems finding work. Turning to men, 78 percent of those who had not worked in over a year said they did not want a job. That included one-third who cited illness or disability, 27 percent who simply had no desire to work, 9 percent who said they were retired, and 5 percent who were in school. Another 4.5 percent were ambivalent about whether they wanted a job, leaving 18 percent who did want a job. But even among men just 6 percent both wanted a job and cited problems finding work. Excluding persons without earnings, then, seems better than including them, unless these figures have changed significantly since 1993. That said, my data still include some people who ended or began spells of labor force nonparticipation that lasted twelve months or more but covered parts of two consecutive calendar years, giving them positive earnings in one or both years. Ideally, if the labor market’s performance is the primary concern, the logic of excluding non-earners calls for omitting these individuals as well.34

I attempt to minimize the number of retirees and students by restricting my sample to persons between the ages of 21 and 60 during the survey (which makes them 20 to 59 years old during at least part of the calendar year for which income is reported). I chose this range as a compromise between being more inclusive and keeping sample sizes large in the PSID on the one hand, and wanting to exclude students and retirees on the other. Since labor force status is reported for the current year while earnings are

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34 I attempted to test the sensitivity of the results for male labor income to the exclusion of long-term non-earners by excluding men who were out of the labor force for twelve months or more (within a year or across years) for the years in which such information is available in the PSID (1985 to 1996). The results were unaffected.
reported for the previous year, excluding students and retirees directly is problematic in the recent survey years, which were conducted only biannually. Furthermore, wives’ labor force status is unavailable in earlier years of the PSID. However, I ran the variance decomposition models excluding people who were students or retired both in the current survey and two years prior to the survey, bracketing the year for which earnings are reported. I also ran all variance decomposition models estimating transitory variances for individuals 18 to 64 years old. The results were negligibly affected.\textsuperscript{35}

I adjust all earnings for inflation using the CPI-U-RS, linking it to the CPI-U for earlier years.\textsuperscript{36}

Finally, the researcher must consider the scale of the measurements used in presenting volatility estimates. It makes little sense, for purposes of describing volatility trends, to present estimates in units of logged earnings squared, which is what variances measure. This is particularly true to the extent that one describes changes in percentage terms. In describing their income variance trends, for instance, Hacker and Jacobs (2008) write,

Although the precise magnitude of the increase depends on the approach to measuring income variance that is used, we estimate that short-term family income variance essentially doubled from 1969-2004. (p. 2)

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\textsuperscript{35} To address the changes in recontact rules and sample definitions, I also ran my variance decomposition results again restricting the sample to exclude the added respondents. Once again, the results were insensitive to this alteration. Additional sample restrictions used in all the results below include requiring individuals to be present in a household in the survey year (not necessarily the same household when comparing multiple years) and requiring them to be PSID sample members. I also require an individual to be a family head (or a head or partner) in all relevant years. For the male head wage volatility trends, I require sample members to have positive hours worked and positive weeks worked. I estimated variance decomposition models in which I held PSID sample membership constant (to address the 1990s changes in the definition), but the results were not much affected.

\textsuperscript{36} See \url{http://www.census.gov/hhes/www/income/income07/AA-CPI-U-RS.pdf}. I also ran variance decomposition models using the CPI-U and ran all models using the CPI-U-RS linked to the CPI-U-X1, and the results were similar.
They note later that the increase in the standard deviation was about 40 percent rather than 99 percent. Instead of a more-than-150 percent increase between 1973 and 1993, they would have shown an increase of roughly 65 percent. Nor do they present their education-stratified estimates or their earnings volatility estimates in standard deviations. I return all of my variances to the original units of logged dollars by taking the square root of the variances. While I present these results as logged dollars in all charts, I generally describe the results in terms of percentage changes in the text to make them more readily interpretable.

Nichols and Zimmerman (2008) note that different decisions about trimming and other methods affect measured volatility levels, which can affect year-to-year changes expressed in percentage terms. None of my trimming decisions or other sample restrictions were made with this issue in mind.

As a last note on presentation, all of my charts are scaled so that the levels of earnings change for similar measures may be visually compared (e.g., downward and upward mobility). Furthermore, the earnings results here and the income results in Chapter Three are consistently scaled in order to facilitate comparing changes across chapters. This presentation decision leads to a lot of empty space in some charts.

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37 Other presentation choices by Hacker and Jacobs also have the effect of exaggerating the increase they find, in particular the scaling of their vertical axis (which obscures the fact that in theory the “cumulative growth” of the transitory variance from its 1973 level could be negative) and (to a lesser extent) their compression of two-year changes from 1996 to 2004 into the horizontal space used for one-year changes from 1973 to 1996 (see their “Figure A”). I discuss these and other problematic presentation decisions by Hacker in Chapter Five.
Measures of Volatility and Instability

Below I present trends in different measures of earnings instability and volatility used in previous work. I also show trends in “pivot volatility”, a new measure that is intended to better capture the concept of volatility. I first show estimates of directional (downward and upward) short-term mobility (relative and absolute). I then present non-directional mobility estimates, including the association of earnings in different years. Next, I present a number of trends based on measurements of dispersion—the dispersion of earnings changes, within-person dispersion, and dispersion of transitory earnings shocks. Finally, I show trends in pivot volatility.

Relative Mobility. I first examine the probability of rising into another quintile or falling from one over two years. I compute quintiles of positive trimmed earnings for each sample and each earnings measure, in every year. That is, quintiles are constructed after excluding persons outside the sample of interest (e.g., persons under age 20, persons with no earnings, men when the sample consists of women) but without regard to whether a person’s earnings are observed in any other year (e.g., the second year over which mobility is measured when constructing quintiles for the first year). This approach means that quintile definitions are not consistent across my analyses, and it also means that upward mobility from a quintile need not be matched by someone moving downward into it. If sample attrition is concentrated among people with relatively high or low earnings, then my approach may indicate more mobility than actually exists. On the other hand, requiring quintile definitions to be based on adults present in both years

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38 A window of two years is the smallest that can be considered over the entire course of the PSID because the survey switched to biennial interviews after 1997.
would mean that only changes in rank relative to a person’s original cohort “count” as mobility, rather than changes in rank relative to the entire labor force in a given year.

**Absolute Mobility.** I compute two measures of absolute mobility: the probability of experiencing a drop or gain of 25 percent or more over two years, and the probability of experiencing a drop or gain of $10,000 or more (in 2007 dollars) over two years. As with the relative mobility estimates, I do not log earnings for these analyses. Because I exclude those with non-positive earnings, I avoid the methodological problem of how to code a change from $0 to a positive amount in percentage terms. Because I trim the bottom and top two percent of observations, I avoid coding small increases from very low initial earnings as large percent changes (e.g., a doubling of earnings from $100 to $200).

**Non-directional Mobility.** I proceed from the directional mobility results to trends in the likelihood of moving up or down (in terms of quintiles or in terms of a 25 percent change in absolute earnings). I also include two measures of intertemporal earnings association: the Pearson product-moment correlation coefficient of logged earnings separated by two years, and Spearman’s rank correlation coefficient.

**Dispersion of Earnings Changes.** Following several previous studies, I compute the standard deviation of two-year changes in logged earnings. This operationalization of instability essentially extends the absolute mobility analysis by looking at the full distribution of earnings changes. I use the difference in logged earnings rather than looking at percent changes because of the asymmetry involved in converting earnings increases and drops of the same dollar amounts to percentages. Imagine a group of three workers making $20,000, $30,000, and $40,000 in one year who each see their earnings rise by $10,000 the next year. The percent changes are 50%, 33%, and 25%. The next
year their earnings fall by $10,000 each – constituting percent changes of -33%, -25%, and -20%. The standard deviation of percent changes has declined, even though earnings simply returned to their original levels.\textsuperscript{39}

**Within-Person Earnings Dispersion.** For every sample member, I estimate the standard deviation of their logged earnings in a nine-year window, using the first, third, fifth, seventh, and ninth years in the window. I then use the mean of these individual standard deviations as a summary of volatility for the year on which the nine-year window is centered.\textsuperscript{40} The earnings observations are for every other year because the PSID switched to biennial surveys after 1997. Like Gosselin and Zimmerman (2008), I require a person to have earnings in at least three out of the five years. This measure is related to the standard deviation of two-year earnings changes. It can be shown that the mean within-person variance using only years $t-2$ and $t$ is equal to the variance of the two-year change multiplied by $(n-1)/2n$, where $n$ is the number of individuals.\textsuperscript{41}

**Dispersion of Transitory Earnings Shocks.** I include two sets of trend estimates for the standard deviation of transitory earnings. Researchers estimating trends in transitory variances or in the variances of innovations to transitory and permanent earnings assume that their results reflect changes in the frequency and magnitudes of these fluctuations, not changes in the long-term trajectories. These types of measures are model-driven estimates and depend on assumptions about earnings dynamics that only

\textsuperscript{39} Dynan et al. (2008) and Dahl et al. (2008) address this problem by using the average of the two years as the denominator in computing percent changes, though Dahl et al. also present results in an appendix using my approach.

\textsuperscript{40} Specifically, I follow Gosselin and Zimmerman (2008) in using earnings in years $t-4$, $t-2$, $t$, $t+2$, and $t+4$ and report volatility in year $t$ as the mean standard deviation across these years.

\textsuperscript{41} Nichols and Zimmerman (2008).
hold imperfectly at best. While these models are indispensable in other research contexts, they are less useful than valid and reliable direct measures for assessing how individuals experience earnings fluctuations.

One of the simplest ways of modeling earnings dynamics is the variance decomposition model of Gottschalk and Moffitt, which is the basis for my first set of transitory dispersion estimates.\textsuperscript{42} Gottschalk and Moffitt begin with a standard model in which an individual’s log earnings at a given time, \( y_{it} \), are decomposed into a permanent component \( \mu_i \) and a transitory component \( \nu_{it} \), with variances \( \sigma^2_\mu \) and \( \sigma^2_{\nu_t} \). Consider the multiplicative model,

\[ z_{it} = \pi_i \omega_{it}, \quad (1) \]

where \( \pi_i \) is (an unlogged) permanent component and \( \omega_{it} \) is (an unlogged) transitory component. Taking the log of both sides leads to the transformed equation

\[ \log(z_{it}) = \log(\pi_i \omega_{it}) = \log(\pi_i) + \log(\omega_{it}), \quad (2) \]

or simply

\[ y_{it} = \mu_i + \nu_{it}. \quad (3) \]

Taking the variance of both sides gives the equation

\[ Var(y_{it}) = Var(\mu_i) + Var(\nu_{it}) + Cov(\mu_i, \nu_{it}). \quad (4) \]

If the permanent and transitory components are uncorrelated, the last term drops out and the variance of income reduces to the sum of the variances of the permanent and transitory components. Furthermore, the covariance of earnings measured in two different years is

\[
\begin{align*}
\text{Cov}(y_{it}, y_{it'}) &= \text{Cov}(\mu_i + \nu_{it}, \mu_i + \nu_{it'}) = \\
\text{Cov}(\mu_i, \mu_i) + \text{Cov}(\mu_i, \nu_{it}) + \text{Cov}(\nu_{it}, \mu_i) + \text{Cov}(\nu_{it}, \nu_{it'}) = \\
\text{Var}(\mu_i) + \text{Cov}(\nu_{it}, \nu_{it'}). \\
\end{align*}
\] 

(5)

If there is minimal autocorrelation of transitory components over long enough periods of time, then if one chooses two years that are sufficiently far apart the covariance in Equation 5 reduces to the first term – the variance of the permanent component. One can then subtract the permanent variance from the income variance in each year to get the transitory variances. It should be reiterated that this model of earnings dynamics is exceedingly simple. In the idealized case in which there is no sample entry or exit, changes in the variance of earnings over time can only arise from changes in volatility.

As noted in Appendix One, Gottschalk and Moffitt actually have come to the view that this model is inadequate and instead emphasize the results from their more complicated error components model.\(^4\) Estimating that model requires using generalized method of moments or nonlinear least squares techniques to produce earnings variance and covariance estimates that best fit the actual earnings covariance matrix observed in the data. The method essentially uses the elements of the covariance matrix as observations and attributes of the covariances as variables used to predict the covariances in a second model. Different assumptions about the dynamics of earnings and the relationships between their components imply restrictions on the model that define its equations and parameters.

\(^4\) Gottschalk and Moffitt (2006), Gottschalk and Moffitt (2007), Moffitt and Gottschalk (2008). See also Shin and Solon (2009), who show how the transitory variance estimates produced by the Gottschalk-Moffitt model will be biased if the returns to permanent individual characteristics are changing over time.
The complexity of error components modeling raises the question of whether it produces estimates that sufficiently improve on those produced by the more basic variance decomposition model. Moffitt and Gottschalk (2008) present comparable results using both methods. They find disparate trends for transitory earnings variances using the two approaches. They argue that the error components model results are superior because the variance decomposition model systematically overstates the increase in permanent earnings variance, thereby understating the increase in transitory earnings variances. If the returns to permanent individual characteristics (estimated in the error components model as part of permanent earnings) are increasing but decelerating, then trends in the covariance of measured earnings will overstate the true increase in the variance of permanent earnings. Essentially, they argue that past returns to permanent characteristics affect future permanent earnings variances, which the variance decomposition model does not allow. Nevertheless, it is possible that the failure of the estimates from their different models to align more closely is due to the PSID data issues described above not being adequately addressed.

Because key volatility research in recent years has relied on Gottschalk’s and Moffitt’s variance decomposition model, I include estimates based on the model here in order to compare them to results using other volatility measures.

There are several important considerations when estimating Gottschalk and Moffitt’s variance decomposition model. First is the lag to use between years in computing covariances. Because the PSID has been conducted biennially since 1997, one is restricted to lags of an even number of years. I present results using four-year lags. The results are not greatly affected, however, if one uses an 8-year lag. Moffitt and
Gottschalk (1995) presented evidence that covariance terms do not change much as one increases the lag, so long as it is of three or four years or more. Moffitt and Gottschalk (2008), however, showed differences in covariances if one uses a ten-year lag rather than a six-year lag. Acs et al. (2007) found that about three in five families experiencing a one-month drop in income of 50 percent or more fully recover within a year, while seven in ten recover 75 percent of their income within a year. Shin and Solon (2009) showed that if the returns to permanent individual traits vary over time, then the risk of biased estimates grows as the lag between years in the covariance term increases. This conclusion, however, hinges on the transitory component of earnings having no serial correlation. If serial correlation is present, then one should use longer lags.

Of course, one could use four-year leads instead of lags. The primary reason not to is that volatility estimates for 2002 and 2004 can no longer be computed, since the most recent currently available PSID survey is from 2005, and it measures earnings in 2004. But because of entry into and exit from the sample, even if the Gottschalk-Moffitt variance decomposition model accurately describes reality, using lags or leads will yield different results.

Consider the ways in practice that earnings variances can change from year to year under the model when using a four-year lag for the covariance term. A change in volatility between years $t-1$ and $t$ can be caused by changes in the transitory variance or by exits from the sample between the two years, but it can also be caused by entries into the sample that occurred in year $t-4$. Workers entering the sample in year $t-4$ will not be included in the computation of the covariance or variance terms in year $t-1$ but will be
included in the computation in year $t$.\footnote{This assumes that one computes the total variance term using the same individuals for which the covariance term is estimated, as I do. That is, to be included in the variance computation, a person must also meet sample restrictions four years earlier. In practice, that means that my estimates are for persons age 24 to 59 rather than 20 to 59.} Entries into the sample in year $t$ will have no effect on the change in volatility and will, in fact, have no effect on volatility trends until year $t+4$.

This is a problem not only because young workers are continuously replenishing the labor force, but because of the PSID recontact efforts in the 1990s, which re-introduced individuals into the sample. If one computes covariance terms using a lead of four years rather than a lag of four years, then both exits and entries from the sample have an immediate effect on volatility trends. Of course, requiring individuals to have earning four years into the future means that the estimates will be for persons age 20-55 instead of 24-59, but at least the inconsistencies introduced by the recontact efforts will be mitigated. I note below the results when leads are used rather than lags where relevant.

Finally, because earnings rise as individuals gain experience and skills, the effect of age on earnings should be adjusted out in the variance decomposition models before estimating variances and covariances. To do so, I pool the PSID waves and regress trimmed logged earnings on a quartic in age, year indicators, and individual fixed effects (stratified by sex in analyses that combine men and women). I use the age coefficients to age-residualize earnings for all analyses. I have also rerun all variance decomposition analyses without the age-residualization, and the results are minimally affected.\footnote{My age-residualization strategy follows Gottschalk and Moffitt (1994). Unlike Gottschalk and Moffitt (2006), I do not residualize income by regressing on education categories (stratified by age and year), not wanting to control for changes in education levels over time. I do not age-residualize earnings for volatility measures other than the transitory earnings dispersion measures.}
The second measure of transitory dispersion that I use is based on an error components model similar to that used by Haider (2001). I model earnings as a function of age, individual fixed effects, a random growth component (individual-specific slopes), and a time-varying transitory component that follows an ARMA (1,1) process:

\[
y_{it} = f(a_{it}, t) + \mu_i + \gamma t + \nu_{it}
\]
\[
\nu_{it} = \rho \nu_{it-1} + \theta \epsilon_{it-1} + \epsilon_{it}
\] (6)

This model implies that once earnings are age- and year-residualized, which removes the first term on the right-hand side in the equation for \(y_{it}\), the covariance of earnings in any two years \(t\) and \(s\) is given by

\[
\text{Cov}(y_{it}, y_{is}) = \text{Var}(\mu_i) + ts \text{Var}(\gamma_i) + (t + s) \text{Cov}(\mu_i, \gamma_i) + \text{Var}(\epsilon_i) \rho^{s-t} [1 + \theta / \rho + (\rho + \theta)^2 / (1 - \rho^2)]
\] (7)

This model may be estimated by stacking the elements \(m\) of the empirical covariance matrix (including the diagonal and above), creating variables \(t\) and \(s\) to designate rows and columns, and using nonlinear least squares to estimate the model

\[
m_c = b_0 + ts b_1 + (t + s) b_2 + \{b_3^{s-t} [1 + b_4 / b_3 + (b_3 + b_4)^2 / (1 - b_3^2)]\} (b_3 + \sum_{j=2}^c d_{cj} b_{j+4})
\] (8)

where the vector of dummy variables \(d\) indicates whether \(t=j\). The coefficients \(b_5\) to \(b_{C+4}\) provide the estimates of each year’s transitory variance, of which I then take the square

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46 I am indebted to Lorenzo Cappellari for sharing his STATA code with me, without which I could not have produced the error components trends below. I require sample members to have at least five non-missing observations in odd-numbered survey years from the time they first appear in the data, and I require them to have no more than 20 percent of observations in odd-numbered survey years missing from the time they first appear. Note that these criteria exclude people who enter the PSID sample in 1999 or later, since they cannot have five non-missing observations.

47 Technically, I regress trimmed logged earnings on a quartic in age and year dummies. The regressions are also stratified on sex for the estimates that combine men and women.
This model is more sophisticated than some, less sophisticated than others, and I have not conducted tests to compare its fit to that of other possible models. I include these results simply to compare volatility estimates that similar models produce to estimates based on other measures of volatility and instability.

**Pivot Volatility.** My last set of estimates involves a new measure I developed to better capture the concept of volatility. A weakness of the volatility and instability measures discussed up to now is that they cannot differentiate between two very different kinds of earnings change: when individuals experience continuously increasing or decreasing earnings, and when rises and falls in earnings follow one another. Increases over time in the measures described thus far may be due to more volatility (the latter kind of earnings change) or due to individuals having an increasing chance of downward or upward mobility (the former kind).

To better capture the concept of volatility, I constructed a measure that is related to the measure of within-person dispersion described above. As with that measure, I focus on changes in individual earnings within a nine-year window centered on the year of interest. Once again, because of the biennial administration of the surveys after 1997, I focus on years \( t-4, t-2, t, t+2, \) and \( t+4 \) within the window. The basis for my measure is the concept of a “pivot”, which I define as a year in which the change in earnings from the previous year is in the opposite direction from the change in the following year. In other words, a pivot year is a year where the direction of change reverses.

For the five years in my nine-year window, there are three potential pivot years—years \( t-2, t, \) and \( t+2 \). When one of these years is a pivot year for an individual, I compute

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48 My model differs from that of Haider in that Haider allows the price of fixed effects and slopes to vary over time.
the absolute value of the percentage change on each side of the pivot year (e.g., for year $t-2$, I compute the absolute value of the percentage change from $t-4$ to $t-2$ and from $t-2$ to $t$). I then average these two absolute values. When no pivot occurs in a potential pivot year, the average is zero. Each individual, then, has three pivot values, corresponding to the three pivot years. A person’s pivot volatility is simply the average of the three values (including any zeroes). I then show trends in the mean pivot volatility across people.\footnote{I require sample members to have non-missing earnings in all five years in the nine-year window.}

This measure has several attractive features. It is readily interpreted as the extent to which the average person’s earnings “jump around” within a window of time. Individuals who experience relatively dramatic turns in their fortunes—experiencing, say, a large income drop after a large rise in income—will have higher pivot volatility than those experiencing a large drop after a small rise in income, who will have higher pivot volatility than those experiencing a small drop after a small rise. Individuals who experience multiple turns in their fortunes will have higher pivot volatility than those who experience a single pivot. Furthermore, the measure does not rely on mathematical concepts unfamiliar to non-researchers, such as logged earnings and standard deviations. It is based solely on percent changes and averages.\footnote{In computing percent changes in earnings over two years, I divide the difference between the two earnings by their average. This ensures that increases and decreases in earnings are treated symmetrically (see the discussion under “Dispersion of Earnings Changes” above). Taking the average of these two “percent changes” is equivalent to summing the two quantities, (difference/sum).}

The estimates I show may be roughly interpreted as the average across people of the average pre- and post-pivot percent change in earnings across possible pivot years. As awkward as this formulation is, it provides a more straightforward description of the magnitude of volatility changes over time than other measures—it allows one to readily
see how big (in percentage terms) typical earnings reversals are today compared with the past.

On the other hand, the measure has some weaknesses. It is based on an arbitrary window of time, and ideally, one would not have to consider only alternating years. Furthermore, individuals must have valid earnings in each of the five years within the window, which—when combined with my trimming—means that they cannot be in the top or bottom two percent of earnings in any of the five years. Nevertheless, the measure more precisely distinguishes volatility from steady earnings change than any of the measures previously used.

Results

Downward Short-Term Mobility

I begin by estimating downward relative and absolute mobility from the PSID to compare them with the results of previous studies. Figure 2.2 shows trends in the percentage of family heads and their partners experiencing three kinds of earnings declines over a two-year period—a fall in the earnings quintile in which one is located, an earnings drop of over 25 percent, and a loss of over $10,000 in earnings. “Earnings” consists of wage and salary income, tips, commissions, bonuses, overtime pay, and self-employment earnings.\(^{51}\)

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\(^{51}\) Technically, I use the PSID measure of “labor income”, which includes the labor part of income from unincorporated businesses, from farming, and from market gardening. Prior to 1993, it also includes the labor part of income from roomers and boarders, and from 1994 onward it includes the asset part of income from market gardening. As noted in the methods discussion, the results are unaffected by these changes in the definition over time.
The charts show that while downward relative mobility does not show much of a trend, the two measures of absolute downward mobility follow the same pattern, which is largely countercyclical. Averaging the figures for 2002 and 2004, roughly one in six prime-age workers was downwardly mobile by each of the measures. Compared with Kopczuk et al.’s (2007) downward mobility results (which give the percent of all workers—head, partner, or other—falling by any amount out of the top two quintiles), my relative mobility results are more volatile, but neither set of results shows much change in relative downward mobility. I find that relative downward mobility was 1 point lower in the early 2000s than in the early 1970s (I compare the average of 2002 and 2004 with the average of 1970 and 1972, pairs of years that bracket the cyclical unemployment peak year).
The trends in the estimated likelihood of an earnings drop greater than 25 percent shown in Figure 2.2 may be compared against the CBO’s results using Social Security records. Overall, the trends match up reasonably well, especially considering that the CBO results include workers who transition from or to $0 in earnings. However, compared with the SSA data, the PSID figures for absolute mobility show more cyclicality, and the figures from 1994-1996 appear too high, perhaps because of methodological changes in the PSID. I show downward mobility defined in this way to have increased about 4 points between the early 1970s and the early 2000s, though there is little secular trend after the early 1980s.

One concern about using percent changes to define absolute downward mobility is that the trend could be driven by the bottom of the earnings distribution, which could be disproportionately influenced by part-time workers. If most 25 percent drops in earnings are from, say, $5,000 to $3,750, then it would be inappropriate to draw conclusions about broad societal patterns from the trend in 25 percent drops. However, as shown in Figure 2.2, the trends are similar if we look at the probability of a $10,000-drop in earnings.

Figure 2.3 shows the same trends when the sample is confined to male heads and the definition of earnings is restricted to wage and salary income.\(^{52}\) Figure 2.4 presents trends for labor income and adds in the very small number of male spouses and partners in the PSID.\(^{53}\) Figure 2.5 displays labor income trends for female spouses and partners. These charts show that relative downward mobility has clearly declined among both men

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52 Wage and salary income are available only for family unit heads in the PSID.

53 Trends in labor income volatility for only male heads are similar to the trends for all men.
and women when they are considered separately, notwithstanding a recent up-tick among men. That is consistent with Buchinsky and Hunt’s (1999) results for the 1980s.

**Figure 2.3. Percent of Male Heads Experiencing Two-Year Declines in Wage and Salary Income**
Figure 2.4. Percent of Male Heads and Partners Experiencing Two-Year Declines in Labor Income

Figure 2.5. Percent of Female Heads and Partners Experiencing Two-Year Declines in Labor Income
Among men, absolute downward mobility increased 6 to 8 percentage points from the early 1970s to the early 2000s. That absolute downward mobility rose while relative downward mobility fell may be explained by rising inequality—as earnings spread out over time, the widths of quintiles increases, making movement out of a quintile less likely for a given absolute change in earnings. Among women, however, the probability of a 25 percent drop fell over time.

The findings for both men and women are consistent with the CBO’s results using trimmed SSA data, and the declines in downward mobility among both groups since the early 1980s are consistent with CBO’s estimates using the full SSA dataset (Dahl et al, 2007, 2008). Women, however, experienced an increase in the likelihood of a $10,000 drop. Because the risk of a woman experiencing a 25 percent earnings decline has not increased, the trend is likely a result of women’s earnings increasing over time; the women earn, the smaller a $10,000 drop is in percentage terms. Furthermore, the risk among women of a $10,000 drop in earnings is lower than other risks of income decline (and generally lower than the risks among men).

As in the SSA data, the gap between the risk that a woman will experience a 25 percent drop and the risk that a man will do so largely closes over time. Unlike the results of Dynan et al. (2007) for PSID family heads in general and Hacker (2007) for male labor income, my results for male heads’ wages and for male labor income do not show an upward secular trend after the early 1980s, though my threshold for experiencing downward mobility differs from theirs. Among men, at least, the PSID figures for 1994-1996 again appear too high, especially in the case of male heads’ wages. This anomaly
may be related to the relatively low variance of men’s wages in the PSID relative to the CPS for those years (see Figure 2.1).

For comparison, Figures 2.6 and 2.7 provide trends in the likelihood of experiencing earnings gains for, respectively, male and female labor income.\textsuperscript{54} Relative and absolute upward mobility have not changed much among men and have declined among women. The exception is that for women, the likelihood of an absolute gain of over $10,000 again increases steadily. Upward mobility is generally more common than downward mobility.

\textbf{Figure 2.6. Percent of Male Heads and Partners Experiencing Two-Year Increases in Labor Income}

\begin{figure}[h]
\centering
\includegraphics[width=\textwidth]{figure2.6.png}
\caption{Percent of Male Heads and Partners Experiencing Two-Year Increases in Labor Income}
\end{figure}

\textsuperscript{54} Results for men and women combined and for male heads’ wages are available from the author upon request.
Summarizing these directional short-term mobility results, the risk of an earnings drop of $10,000 or more has risen among both men and women. Among men, this has also translated into a rise in the risk that earnings will fall by 25 percent or more, but because inequality has grown and because the trend in large earnings losses is so cyclical, the probability of falling one or more quintiles in the earnings distribution has actually declined. Among women, the rising risk of a $10,000 drop has coincided with a decline in the risk of a 25 percent earnings loss, implying that the growth of women’s earnings accounts for the former trend. If what matters for earnings drops is the size of percent changes, risk has increased among men and fallen among women since the early 1970s. However, there has been little secular change in risk for either since the recession of the early 1980s. The greater risk of a drop among men during the 1970s and early 1980s was
not accompanied by a greater likelihood of large earnings gains. Thus far, then, the
evidence suggests that any great risk shift was confined to men and to the 1970s and early
1980s.

The patterns of downward short-term mobility are reasonably consistent across
measures and datasets, at least when my PSID estimates are compared with earlier results
from other datasets. Where inconsistencies exist across studies, they are apparently due
to minimal amount of change in downward mobility after the early-1980s recession.

Non-Directional Short-Term Mobility

Is it possible to reconcile estimates for other measures that more directly
correspond to non-directional instability and volatility? The four measures of non-
directional short-term mobility in Figures 2.8 to 2.11 provide a bridge between the
directional mobility results above and the dispersion estimates to follow.

Probability of Earnings Change in Either Direction. Figure 2.8 displays
trends for all heads and partners in non-directional short-term mobility. The top two lines
show trends in the probability of moving up or down one or more quintiles and moving
up or down 25 percent. Figure 2.8 clearly shows that there has been little change in the
probabilities since the early 1970s, with both hovering around 40 percent. That is in
contrast to the results of Kopczuk et al., which show an increase in the probability of
falling from the top or rising from the bottom in the early 1970s, followed by a decline
thereafter.
Figure 2.8. Non-Directional Labor Income Mobility Trends

Figure 2.9. Non-Directional Male Heads’ Wage Mobility Trends
Figure 2.10. Non-Directional Male Labor Income Volatility Trends

Figure 2.11. Non-Directional Female Labor Income Volatility Trends
Figures 2.9 and 2.10 show the same trends for male wages and labor income, respectively. They indicate that among men non-directional relative mobility was flat or declining over the period while non-directional absolute mobility increased by roughly 5 percentage points—all of the increase occurring after the early-1990s recession.

Figure 2.11 shows that among women, non-directional mobility in either a relative or absolute sense declined over time. This result is unsurprising, given that Figures 2.5 and 2.7 showed declining absolute mobility in both directions among women.

The relative mobility trends for men and women are largely consistent with other studies, and again, where the previous research was contradictory (for the 1970s), the trend I find is flat. The past research on non-directional absolute mobility is sparse, but my results contradict those of Dynan et al., and my results for men contradict those of Dahl et al.

**Intertemporal Earnings Association.** The bottom two lines in Figure 2.8 show trends in two correlation coefficients—the Pearson and Spearman coefficients—each subtracted from one, corresponding to intertemporal association of logged earnings and earnings ranks. The two measures follow similar trends, indicating rising non-directional mobility, particularly during the late 1980s and early 1990s. Figures 2.10 and 2.11 reveal that the trends for men and women differ once again, with men seeing rising mobility (since 1989) and women declining mobility.

Figure 2.9 reveals that the rise among men disappears if one looks only at wage and salary earnings, revealing a difference that will recur in the dispersion-based measures below. Wage and salary mobility and labor income mobility take markedly different trajectories after 1990. These results imply that self-employment earnings are
largely responsible for the apparent rise in mobility among men in Figure 2.10.\textsuperscript{55} The literature review of previous research in Appendix One found little consistency against which to compare my results.

Overall, then, my estimates of non-directional short-term mobility indicate that it rose among men and declined among women. The likelihood of a large earnings change among men increased only after the recession of the early 1990s. Intertemporal association of male earnings has also increased since 1990, apparently driven by changes related to self-employment.

\textit{Dispersion of Earnings}

The next set of measures I examine is based on dispersion of earnings or earnings changes. They correspond more closely to the concept of volatility than the mobility measures in the sense that they tend to focus on entire distributions of earnings changes across or within individuals or restrict themselves to components of earnings that are theorized to be transitory. While each of them has weaknesses as a measure of volatility, consistent results across different conceptualizations of volatility would suggest that my trend estimates are meaningful.

\textsuperscript{55} The labor income measure differs from the wage and salary measure in that it includes a very small number of male “non-heads”, but such men are too rare in the PSID to account for the difference. The labor income measure also includes other sources of earnings besides self-employment income, including bonuses, overtime, tips, and commissions. Alternatively, the increase after 1990 could be due to changes in the way that men’s labor income was measured over time in the PSID. That variable was an aggregation of up to six separate reported income amounts in the 1968 survey, up to seven during the 1980s and early 1990s, and up to eleven or twelve thereafter. The variable for men’s wages and salary income was an aggregation of no more than two income amounts during any of the PSID survey years. Not only does this difference between the two measures make for greater inconsistency in the labor income measure, but many of the changes in the PSID administration discussed above—from topcoding to question structure—had bigger effects on the labor income component variables than the wage and salary variable. All that said, it is unclear why these changes should produce \textit{steadily} declining correlations (steadily increasing mobility).
Dispersion of Earnings Changes. The upper line in Figure 2.12 measures the trend in the standard deviation of two-year changes in logged labor income. The measure shows little change over time, though a cyclical pattern is evident, with increases in volatility during most recessions. Volatility increased by just 3 percent over the thirty years between the early 1970s and the early 2000s. The pattern somewhat resembles the Dahl et al. results looking at the standard deviation of either one-year percent changes or one-year differences in logged earnings, but the resemblance would be closer if my early-1980s figures were higher, which would produce a steeper decline in volatility during the 1980s and an overall decline from the early 1980s to the early 2000s. However, the decline in the Dahl et al. study was largely due to a drop in the number of workers who reported $0 in earnings in one of the two years over which earnings change was measured. Since my results do not include persons without earnings, it is reasonable to think that Dahl et al.’s SSA figures would match mine if we treated persons without earnings in the same way. My slight increase over the entire span of years is also consistent with Dynan et al.’s (2008) finding of a slight decline among all heads and partners. The bottom line is that overall, volatility has not changed much in the past thirty years by this measure.

Disaggregating men and women, however, again leads to a different conclusion. Figure 2.13 and Figure 2.14 show trends in this same measure for male heads’ wages and male labor income, respectively, indicating increases of roughly 20 percent and 60 percent from the early 1970s to the early 2000s. The trend in Figure 2.13 is very close to that found by Shin and Solon using the same dataset and a very similar measure, though their estimates fluctuate more because the results are shown as variances rather than
standard deviations. My estimates are also qualitatively similar to those of Moffitt and Gottschalk (2008).

Between the early 1980s and 2004, Figure 2.13 shows a basically flat trend. Even though my PSID results include only heads, the results match reasonably well with the small decline found in the CBO results among men, especially if one factors in the effect of including $0 reports on the CBO findings. Figure 2.14, which includes self-employment income (unlike the CBO’s SSA results), shows an increase from the early 1980s to 2004. It is unclear how to reconcile this increase with the CBO’s results. The CBO found that when it included self-employment earnings after 1990, the trend in
earnings volatility was unaffected. On the other hand, Dynan et al.’s PSID-based research indicates that when the self-employed are included, male earnings volatility increases notably. Their 2007 paper showed a 40 percent increase in male heads’ labor income volatility that dropped to a 20 percent increase when those without a business interest were excluded. Their 2008 paper found a 70 percent increase in male labor income volatility between the early 1970s and the early 2000s, which is consistent with my findings.

My results are also consistent with Dynarski and Gruber’s (1997) finding that volatility rose by about a third from 1970 to 1991, though I do not find the large increase

56 Dahl et al. (2008). As my Figures 2.3 and 2.4 show, including self-employment does not alter my earnings-drop results, even though it does alter my dispersion-of-changes results (unlike those of ČBO).
in volatility during the late 1980s that they show. Cameron and Tracy (1998) show a 17 percent increase from 1972 to 1996 (when measured in standard deviations). I show a 44 percent increase, though my trend for the 1970s is flatter than theirs, and I find a bigger increase in the early 1990s. My results are also broadly consistent with those of Abowd and Card (1989) and of Baker (1997). To summarize these results, it appears that volatility among men increased between the early 1970s and early 2000s, by about 20 to 30 percent when excluding self-employment earnings, and largely during the early-1980s. Growth in volatility was two to three times as large when self-employment earnings are included, with the difference entirely due to trends since the early 1980s.

Figure 2.14. Male Labor Income Dispersion and Pivot Volatility Trends
Finally, I present results for women in Figure 2.15, which reveal that the overall flatness of the volatility trend in Figure 2.12 is the result of declining volatility among women balancing out rising volatility among men. My results align well with the CBO results for women, showing a 21 percent decline from 1981 to 2002, compared with the 20 percent decline in the SSA. The decline I find from the early 1970s to the early 2000s also closely matches that of Dynan et al. (2008).57

![Figure 2.15. Female Labor Income Dispersion and Pivot Volatility Trends](image)

I also estimated the mean absolute deviation (MAD) of two-year changes in logged earnings. The MAD is computed by obtaining the deviations of two-year earnings changes from the mean two-year earnings change, and then taking the average deviation. This measure of dispersion, unlike the standard deviation, does not weight large deviations more heavily than small ones in summarizing the “typical” deviation from the central tendency. The results, however, were generally unaffected. See Gorard (2005) for a critique of the conventional use of the standard deviation.

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**Within-Person Earnings Dispersion.** Returning to Figure 2.12, I now examine trends in volatility measured as the average within-person standard deviation across a number of years, displayed in the second-lowest line in the chart. The trend shows a 10 percent increase from 1973 to 2000, which fell at roughly comparable points in the business cycle.\(^{58}\) While my increase between 1973 and 1998 is comparable to Gosselin and Zimmerman’s increase of 11 percent in the PSID, my trend along the way looks quite different from theirs, particularly given that our measures are very similar. Gosselin and Zimmerman (2008) find a sharp increase in volatility between 1986 and 1990, followed by an equivalent decline from 1990 to 1998, a pattern that is unique among all the volatility studies. Gosselin and Zimmerman’s SIPP results are more consistent with my PSID estimates over this period than with their own PSID results, though the absence of data points in key years makes the comparison difficult. Gosselin and Zimmerman find a 10 percent increase in volatility between 1983 and 1996 in the SIPP, while I find no change in the PSID.

Between 1973 and 1991, Comin et al. (2006, forthcoming) show a doubling of earnings volatility (measured as an average of within-person variances) in the PSID. That increase is much bigger than the 7 percent increase I show (measured as an average of standard deviations). Unlike Keys (2008), I find similar levels of volatility in the 1980s and 1970s. The flatness of my results matches up well with the estimates of Nichols and Zimmerman (2008) when they take various transformations of earnings.

We can return to Figures 2.13 and 2.14 to consider the trend in male volatility measured as within-individual dispersion. Figure 2.13 shows a 20 percent increase

\(^{58}\) The years for which this measure is available are reduced because it requires observations every other year in a nine-year span, which affects the beginning and end of the time series as well as two years in the 1990s (due to the shift to biennial surveying).
between 1973 and 2000, which is consistent with the roughly 30 percent increase found using the standard deviation of log earnings changes. The 16 percent increase from 1974 to 1986 is an order of magnitude smaller than the 120 percent increase Gottschalk and Moffitt (1994) found for white male heads, though they report volatility in terms of average variances. Like Daly and Duncan (1997), I find higher volatility in the 1980s than the 1970s.

Figure 2.14 includes all labor income. As with the trend in the dispersion of male labor income changes, including income from self-employment produces a steadily increasing trend. Between 1973 and 2000, I show an increase of one-third, consistent with the 40 percent increase in the standard deviation of earnings changes over the same period.

Finally, Figure 2.15 displays the trend for women, which again is similar to the trend for the dispersion of earnings changes. Volatility declines 14 percent from 1973 to 2000, compared with 23 percent for the standard deviation of earnings changes. There are no previous estimates for women using this approach.

**Transitory Earnings Dispersion.** My final set of dispersion-based volatility estimates is comparable to previous research examining trends in the transitory variance of earnings. The second-highest line in Figure 2.12 shows the square root of the transitory variance of labor income when I use Gottschalk and Moffitt’s variance decomposition model. Between the early 1970s and the early 2000s, volatility increased by about 15 percent. The trend is quite similar to that for the dispersion of earnings changes. The bottom line shows the trend in the transitory variance using a more
complex error components model. In contrast to the variance decomposition model, this model indicates that volatility declined by over 20 percent.

Once again these overall patterns obscure significant differences by sex. Figure 2.13 shows trends for male heads’ wage and salary income, and Figure 2.14 shows trends for male labor income volatility. Through 1990, the variance decomposition trends in the two figures track each other very well, and both clearly follow a cyclical pattern even after 1990. Both also track the trends in dispersion of earnings changes. Male heads’ wage volatility increased 67 percent from 1973 to 2004 based on these estimates, while male labor income volatility increased 76 percent from the early 1970s to the early 2000s.

My variance decomposition estimates for male heads depart dramatically from Moffitt and Gottschalk’s latest paper (2008). However, my estimates track Moffitt and Gottschalk’s preferred error components model estimates very closely. The fact that my male variance decomposition estimates and error components estimates track each other well and also track Moffitt and Gottschalk’s error components estimates too calls into question their conclusion that variance decomposition estimates of transitory variance are to be avoided, given that this conclusion stemmed from the lack of congruence between their own variance decomposition estimates and the other estimates they showed.

My results also depart from Gottschalk and Moffitt’s 2006 results, which should be more comparable to mine than those in their 2008 paper. I find a 46 percent increase in heads’ wage volatility from 1974 to 2002, which is notably smaller than their 145 percent increase. The difference is largely due to my taking the square root of the transitory income variances to return them to non-squared units of measurement. Putting the Gottschalk-Moffitt estimates on that scale would show an increase of roughly 55-60
percent.\textsuperscript{59} But aside from the endpoints, the wage volatility trend in Figure 2.13 differs from their results in years between 1990 and 1998. I find a sizable decline in volatility between 1992 and 1993, while Gottschalk and Moffitt report a small increase, and I find a small increase in volatility from 1996 to 1998 while they find a decline. Gottschalk and Moffitt also find a bigger increase between 2000 and 2002 than I do. My variance decomposition estimates also fail to track the SSA estimates that Schwabish estimated for Gottschalk and Moffitt (Gottschalk and Moffitt, 2007).

Comparing Figure 2.14 with the results of Hacker and Jacobs (2008), my 75 percent increase in labor income volatility I find from 1973 to 2004 is very close to the increase they would have found if they had expressed their results in standard deviations.\textsuperscript{60}

The volatility estimates for men using the Gottschalk-Moffitt variance decomposition model show larger secular increases than my other measures. But there is reason to think that these estimates overstate the true increase over time. A glance back at Figure 2.1 reveals that the total earnings variance increases among men after 2000 in the PSID but not the CPS, and my estimates indicate that the increases in the transitory variances of male earnings after 2000 are driven by increases in the total variance of earnings (not shown). Above, I noted that the CATI instrument changed in the 2003 survey, which could have introduced a discontinuity between pre-2002 volatility

\textsuperscript{59} Gottschalk’s and Moffitt’s income residuals come from regressing on education categories (stratified by age and year) rather than on age and year, and they trim the top and bottom 1 percent of the wage and salary distribution rather than the top and bottom 2 percent, among other smaller differences in our methods.

\textsuperscript{60} This is unsurprising given that I shared my statistical programming with Jacobs between the release of Hacker’s initial numbers and their publication of the EPI report.
estimates and those for 2002 and 2004. If one compares the cyclical peak years of 1973
and 2000, the increase in wage or labor income volatility is roughly 55 percent.61

Furthermore, one can get a better estimate of the 1973 to 2000 change by
computing the covariance term using years $t$ and $t+4$ rather than $t-4$. When I re-computed
volatility estimates using 4-year leads, I found that male heads’ wage volatility and male
labor income volatility increased by just 25 to 45 percent, which is similar to my
estimates using other volatility measures. The difference is presumably related to the
problem that leaving and entering the sample poses for variance decompositions based on
covariances with lagged income, which I noted above.

Turning to the error components estimates, I find that transitory wage volatility
declined slightly from the early 1970s to the early 2000s. Transitory earnings volatility
(including self-employment earnings) increased by roughly a third.

In finding declines in transitory variances from the mid-1980s to the mid-1990s,
my error components results depart notably from most earlier research, though it is worth
pointing out that my estimates generally track my estimates using other measures of
volatility. I find an increase of roughly 25 percent from 1973 to 2004, compared with
Moffitt and Gottschalk’s two-thirds increase (Moffitt and Gottschalk, 2008). Haider
found an increase of roughly 60 percent in male labor income volatility (expressed in
standard deviations) between 1971 and 1991, compared with my 27 percent increase,
though my estimates show the same trends as his. Stevens (2001) found a 35 to 40
percent increase in male heads’ wage volatility from 1973 to 1991 (expressed in standard
deviations), while I find a 28 percent increase. Daly and Valletta (2008) found a

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61 Interestingly, the estimates I produced when using the full core sample and applying the survey weights
showed a 50-60 percent increase in wage or labor income volatility from the early 1970s to the early
2000s—that is, incorporating the suspicious 2002 and 2004 data points.
roughly 15 percent increase in volatility from 1979 to 1996 among white male heads; I find a 19 percent increase among all male heads. While Mazumder (2001) finds a 25 percent increase from 1984 to 1997, I find a 15 percent decline from 1984 to 1996.

Finally, as with other volatility measures, Figure 2.15 shows that volatility among working women has declined. The transitory standard deviation based on the variance decomposition model was 12 percent lower in the early 2000s than in the early 1970s. The decline is clearer if the 2004 data point is discounted—25 percent from 1973 to 2000 (peak to peak) if the covariances are computed using four-year leads. The error components model indicates a decline of over 40 percent from the early 1970s to the early 2000s. Once again, there are no previous estimates to compare with these figures.

_Pivot Volatility_

Finally, I present my estimates of pivot volatility—the extent to which individuals experience sizable earnings reversals. These estimates are also shown in Figures 2.12 through 2.15. In general, trends in pivot volatility resemble those for within-person dispersion. The estimates reveal very modest increases in volatility for men. Looking at wage and salary income, the average pre- and post-pivot earnings change over a nine-year window increased from 12.5 percent to 13.8 percent between 1973 and 2000. Including all labor income increases volatility from 14.1 percent to 17.7 percent. Once again, volatility declined among women, from 21.5 percent to 17.9 percent. Overall, then, these estimates reinforce the results above, though they more clearly show how small the increase over time has been.
Discussion and Conclusion

My different measures of earnings instability and volatility yield similar conclusions. Overall, there has been little change in earnings movements since the early 1970s for the typical worker, particularly in the past twenty-five years. However, this stability masks a sizable difference between trends for male and female workers.

For the typical male, downward earnings mobility (measured in terms of a large percentage decline) has increased, but the increase was largely confined to the 1970s and early 1980s. Among men with no self-employment earnings volatility increased by 20 to 30 percent using the dispersion-based estimates that are not model-driven. This increase ended in the early 1980s. The typical earnings reversal as measured by my pivot volatility measure was hardly any larger in 2000 than it was in 1973. If I include self-employment earnings, volatility continued to rise among men after the early 1980s. Volatility including self-employment earnings rose between one-third and 60 percent among men from the early 1970s to the early 2000s. Again, however, the typical earnings reversal grew only modestly.

When self-employment income is excluded my estimates of the trend in volatility for men – increases through the early 1980s followed by a relatively flat trend – is consistent with the CBO estimates using Social Security Administration data (Dahl et al., 2007, 2008). But CBO did not find that adding in self-employment earnings altered their results. Why then, does the PSID tend to show increases in labor income volatility after the early 1980s?

The increase in labor income volatility in the PSID is due to a rise in the volatility of income from self-employment rather than to a rise in the frequency of self-
employment (analyses not shown). Self-employment income in the PSID is allocated somewhat arbitrarily between labor and asset income and between heads and “wives”, but the results were not much different when I allocated all self-employment income to heads’ labor.

While the question remains open, one possibility is that sample attrition has biased PSID estimates over the past 15 years. However, as noted above, previous research on attrition from the PSID has failed to find any evidence of bias. The fact that the PSID did not try to survey a nationally representative sample until 1999 also suggests another possibility. Until 1997, the PSID represented the national population as of 1968 and its descendants. This population excludes nearly all post-1965 immigrants and their descendants. The PSID added immigrant samples in its 1997 and 1999 waves, and the weights were constructed to yield a nationally representative sample moving forward.

Beyond these explanations, it may also be relevant that the labor income variable in the PSID is composed of a number of component variables and that the composition of the overall variable – as well as the measurement and editing of the components – has changed over time. The variable for male heads’ wages, on the other hand, remains nearly completely consistent over time. It is entirely possible that all of the changes in the PSID in the 1990s – recontact efforts, changes in following rules, a switch to more detailed survey questions, modifications to the CATI software, changing income editing procedures, budget cuts, and a more-than-doubling of the interview length – affected the composite labor income measure more than it affected the measure of male heads’ wages. A final alternative is that self-employment earnings are better measured in the PSID than in the Social Security Administration data. Unfortunately there is no obvious way to
determine which of these explanations accounts for the observed difference between the two data sets.

Finally, among women, downward mobility and earnings volatility declined fairly steadily over the period. The first-of-their-kind estimates I provide for trends in women’s within-person dispersion and transitory earnings dispersion are consistent with the results from earlier studies that use different measurements. A reasonable interpretation of these trends is that as women became more attached to the labor force over time, increasing their hours and weeks worked, their earnings became less volatile. In results not shown, I find that these downward trends among women closely track Bureau of Labor Statistic figures on the percentage of women who work less than full time (including those who do not work at all).

My results are broadly consistent with earlier studies, save a few oft-cited papers that find bigger increases using the PSID. The evidence on earnings volatility provides little indication of a recent structural “risk shift” in the economy. Volatility among men increased during the 1970s and during the recession of the early 1980s, but there has been little secular change since then. The earlier increase may constitute a “risk shift”, but its timing does not match recent accounts that argue for a deterioration in workers’ security in the 1990s. Nor does it appear as large as some previous studies have implied. Furthermore, if the SSA short-term mobility estimates of Kopczuk et al. are any indication, volatility levels were probably higher prior to 1960. Earnings movement may have increased in the last twenty-five years among self-employed men. Volatility among women has declined over the period. Finally, if the indicators of risk in Chapter Five are any guide, instability and volatility were probably no worse in 2008 than in 2002 or 2004.
Of course, just because the risk-shift narrative does not appear to fit well in terms of earnings does not mean that it fits poorly in other domains, such as worker benefits or family income. The next chapter estimates trends in family income instability to see what the evidence shows.
Chapter 3

The Very Small Risk Shift: Trends in Family Income Changes

*Our main finding is that family incomes are lot less stable than they used to be. In particular, the chance that families will experience catastrophic drops in their income has risen dramatically over the last generation.*

– Jacob S. Hacker and Elisabeth Jacobs, April 2009.¹

Few academic topics have attracted as much recent attention in political and policy circles as the question of whether family incomes are less stable today than in the past. There are clear reasons why trends in family income volatility and instability might differ from the earnings trends from Chapter Two. First, families are usually aggregations of individuals, some of whom have earnings and some of whom do not.² Income volatility within a family can change not only because the earnings of individual workers become more or less volatile, but because the number of workers changes more often, because their hours of work are interdependent, or, if income is adjusted for the fact that families of different sizes have different economic needs, because the number of non-workers changes. Changes in how closely the economic situations of individual family members resemble each other will also affect income volatility trends. If husbands and wives, for example, work in increasingly similar jobs over time, then to the

¹ Hacker and Jacobs (2009).

² I will generally use “family income” rather than “household income” to describe results. The PSID asks about the incomes of “family unit” members. Family units in the PSID include most cohabiting partners as well as relatives of the family unit head and his cohabiter (“heads” in the PSID are almost always men when the family unit includes a couple). There are only a small number of households in the PSID with multiple family units interviewed. Unlike Census Bureau use of “family”, my use here is intended to include family units consisting of individuals living without relatives.
extent that volatility hits some occupations harder than others, family incomes will be more unstable than when husbands and wives performed less similar work.³

Second, earnings are only one type of income. The streams of income from other sources, such as from investments, family and friends, and public programs, can also become more or less volatile, compensating or exacerbating earnings volatility. Furthermore, changes in the flow of income out of a family – to relatives or friends and in the form of taxes – can affect volatility trends if income is measured net of these transfers.

In this chapter, I provide a range of trend estimates describing income instability and volatility over the past 35 years. The goal throughout is to determine what consistent findings recur and how income volatility trends should be characterized. I find that neither the risk of income drops nor volatility per se increased much from the early 1970s to the early 2000s.

Before presenting my results, however, it is worth providing some context for the boomlet of recent research in this area, and that begins with the influential but flawed work of political scientist Jacob Hacker.

**Jacob Hacker and *The Great Risk Shift***

While not the first to examine trends in income instability, the recent wave of research on the subject undoubtedly began with and proceeded from Hacker’s work. Beginning in early 2004, mostly in outlets aimed at a general audience, Hacker promoted the view that income volatility had increased dramatically in recent years. This was the primary evidence for what he called a “Great Risk Shift” that heralded a world in which

³ Nichols and Zimmerman (2008).
Americans were increasingly vulnerable to economic calamity. Hacker’s very visible findings created an impression among political and policy observers and participants that income volatility constituted a neglected economic problem that indicated and explained widespread economic anxiety. However, the bleak impression Hacker created was the result of a combination of questionable data presentation decisions and inadequate attention to how sensitive his results were to potential data problems.

Hacker’s first published volatility estimates came in a *New York Times* op-ed in January of 2004, in which he announced that according to the Panel Study of Income Dynamics (PSID), “the instability of family incomes was roughly five times greater at its peak in the 1990’s than in 1972.” This finding, shown in a dramatic chart accompanying the piece, was Exhibit A of his argument that “the economy has become more uncertain and anxiety producing for most of us—not just over the past three years, but over the past 30.”

To measure volatility, Hacker used the Gottschalk-Moffitt variance decomposition model discussed in Chapter Two and below. In displaying his results, he presented averages of five years of volatility estimates, centered on the year in question.

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4 Hacker (2004a).

5 In these initial results, Hacker used the covariance of current income and income lagged five years to estimate the permanent income variance. Since the data Hacker relied on only went as far back as 1970, the five-year lag meant that 1974 was the earliest year for which he could estimate a transitory variance (since the 1975 survey measures 1974 incomes). For the years 1976-1994, the figures in the chart are averages of five years’ transitory variances, centered on the transitory variance of the year in question. This raises the question of how it is possible to have data points for 1972-1975 and for 1995-1998, since it is not possible to compute five-year moving averages centered on any of these data points. Though none of the many publications Hacker has written on his estimates includes an answer, I have determined that the figures for 1972-75 are the averages of the *available* transitory variances within a five-year window. For 1975, the four estimates for 1974-1977 are used; for 1972, only the 1974 estimate is used (that is, the 1972 data point is actually the 1974 estimate). The same holds for the figures for 1995-1998, except that because there was no 1998 survey wave, the 1999 survey wave (measuring 1998 income) is treated as part of the window—for 1995, the figure in the chart is the average across the years 1993-96 and 1998; for 1998, the figure in the chart is the average across the years 1995, 1996, and 1998.
Hacker’s estimates showed pre-tax family income volatility increasing from 1972 to the mid-1980s, with a relatively flat trend in the second half of the 1980s and a sharp increase from 1989 to 1994. Volatility then declined sharply from 1994 to 1996 and continued to decline from 1996 to 1998. These estimates are reproduced in Figure 3.1 below.\(^6\)

![Figure 3.1. Jacob Hacker Trends in Family Income Volatility, New York Times Op-Ed, January 2004](http://pantheon.yale.edu/~jhacker/PSID_Data_NYT.htm)

Comparing the 1972 estimate in Hacker’s chart to the peak volatility level in the early 1990s shows volatility 3.7 times higher in the latter period. The basis for stating that volatility was 5 times higher lay in comparing the raw 1974 and 1993 levels; volatility in 1993 was 4.75 times the 1974 level. However, as Figure 3.1 shows, had

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\(^6\) The figures behind these numbers come from Hacker’s personal website at [http://pantheon.yale.edu/~jhacker/PSID_Data_NYT.htm](http://pantheon.yale.edu/~jhacker/PSID_Data_NYT.htm).
Hacker displayed the raw transitory variance levels, rather than five-year moving averages, what would have stood out would have been not just the unprecedented rise in volatility between 1991 and 1993, but the fact that it was completely reversed from 1993 to 1998. Such a dramatic decline would have severely weakened the argument that family income was growing more volatile, raising economic “risk” in America.

The large decline in volatility from its early 1990s peak led Hacker to characterize the overall trend in volatility over the 24 years, alternatively, as having “more than doubled between 1974 and 1998”, as having “become two to three times more unstable in the past three decades”, and as having “never fallen below twice its starting level” after it peaked, all of which were more or less accurate ways of describing the 110 percent increase in the raw transitory variances.\(^7\) As it turns out, however, Hacker’s op-ed and his subsequent estimates greatly exaggerated how volatile volatility was after 1990.

By early 2005, Hacker had added the 2001 PSID to his database, allowing him to produce a volatility estimate for 2000. The raw figure indicated no change since 1998, leaving volatility below its 1991 level. However, Hacker continued presenting his results as moving averages, which continued to show volatility well above its early 1990s levels.\(^8\)

In 2006, Hacker incorporated newly-available data from the 2003 PSID. As quoted in a June 2006 *U.S. News and World Report* article just a few months before his book, *The Great Risk Shift*, was published, Hacker indicated that he had, “just analyzed the 2002 data. Family income volatility increased by 50 percent over the past two years,

\(^7\) Hacker (2004b, 2004c, 2004d).

so it is now three times its early-1970s level.”

With the PSID trend line now indicating that volatility had tripled since 1974 and was once again climbing skyward, Hacker switched from showing moving averages to reporting the raw volatility estimates in *The Great Risk Shift* (henceforth, *GRS*).

Figure 3.2 shows Hacker’s pre-tax family income volatility estimates, as reported in *GRS*. In all, volatility increased by 360 percent from 1974 to 1994, declined by 57 percent from 1994 to 1998, and increased 50 percent from 2000 to 2002. Volatility more than doubled from 1991 to 1993 alone according to these results. Post-tax income volatility showed similar patterns and grew 130 percent over the entire 1974-2002 period.

For comparison, Figure 3.2 also shows what Hacker’s results would have looked like if he had stayed with his prior decision and shown the results as moving averages. Notably, volatility declines below its 1991 level and does not turn upward.

Sticking with Hacker’s presented estimates, however, the large volatility changes after 1990 stand in stark contrast to the smaller year-to-year changes in earlier decades. It was the starkness of the post-1990 increase in volatility that Hacker presented in his *New York Times* op-ed that led me to begin efforts to replicate Hacker’s results prior to the publication of *GRS*. These efforts culminated in a concession from Hacker that his

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9 Pethokoukis (2006).

10 Hacker (2006a). The figures are from a web page that was on Hacker’s personal website for a time but is no longer available. Hacker used the PSID and a version of the PSID called the Cross National Equivalent File that includes modified income variables and tax estimates. His measure of pre-tax family income was a composite variable that included taxable income, public energy subsidies, and the rental value of free housing from the PSID, and public and private transfers from the CNEF. The post-tax measure subtracted federal and state income taxes, and payroll taxes taken from the CNEF, and property taxes taken from the PSID. Hacker adjusted income for family size by dividing by square root of family size, and he then logged the measures and adjusted them for inflation using the International Monetary Fund CPI (from the CNEF). His sample was confined to adults 25-61, including the entire core sample. He did not, however, apply the sample weights. Finally, in computing covariance terms, Hacker used 5-year lags, though to obtain the 2002 estimate, he used a 6-year lag, since there was no 1998 survey to provide 1997 income estimates. These details were obtained through personal communication with Hacker in 2006 and 2007.
figures were inaccurate and publication of a revised paperback edition of *GRS*, after I was able to show that the large post-1990 swings were driven by Hacker’s failure to address year-to-year changes in the number of families reporting very small incomes.\(^{11}\)

I have since updated my replication attempts based on (incomplete) information provided by Hacker. Figure 3.3 again shows Hacker’s pre-tax family income volatility figures from *GRS*, along with my best replication attempt working with the information that I was able to obtain from him.\(^{12}\) While I could not replicate the full extent of the post-1990 swings in *GRS*, my trend line follows the same broad patterns as Hacker’s.

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\(^{11}\) Winship (2007), and personal communication with Hacker (2007). Elisabeth Jacobs served as a useful sounding board and liaison to Hacker in the early months of my explorations, before joining Hacker’s research team.

\(^{12}\) My best replication combines household taxable income, energy subsidies, and free rent from the PSID, and private and public transfer income and windfall income from the CNEF. I have attempted to code
As discussed in the previous chapter, when incomes are transformed by taking their natural logs, volatility estimates relying on the Gottschalk-Moffitt variance decomposition model are strongly influenced by year-to-year changes in the number of very small reported incomes. Taking natural logs drops cases with non-positive incomes, since their natural logs are undefined, and it also makes differences between very small positive incomes very large. I demonstrate this in Figure 3.3 by showing how the volatility trend changes if one either bottom codes all incomes—positive or not—at $4,000 (or $333 per month) before taking natural logs or trims the bottom three percent these variables consistent with incomplete code provided by Hacker in early 2007. My replication also incorporates several uncorrected data problems that may have affected the results. See note 11 for additional methodological details. The replication is imperfect both because Hacker had access to non-public CNEF data for the years 1970-1979 and 2002 that he could not share and because he did not provide me the full code that would have allowed me to replicate his figures based on the 1980-2000 data. He did provide me with enough information to confirm that his trends in taxable family income were incorrect and conceded as much in spring of 2007.
of incomes in each year. Rather than the roughly 150 percent increase in volatility from 1989 to 1993 that Hacker found, or the doubling that my replication attempt shows, I find an increase of about one-fourth using the bottom-coded estimates, or two-thirds using the trimmed estimates.

Hacker’s family income measure does not include retirement income, but even among individuals between the ages of 25 and 61, the retirement income of family members has an important moderating influence. Hacker’s measure also effectively double-counted alimony payments from 1994 onward. Figure 3.3 shows the volatility trend when the basic pre-tax family income measure in the PSID is used (after trimming), which corrects these problems. I find an increase in volatility of 65 percent over the 1974-2002 period, rather than the 200 percent implied by Hacker’s estimates. Similarly, volatility increases by about 75 percent if the sample is confined to the nationally representative subsample of the PSID, which I found to more closely track CPS trends in income variance than the full core sample that Hacker used. This latter trend line also shows a smaller bulge in volatility in the early 1990s.

As discussed in the previous chapter, presenting volatility estimates as variances leaves the scale in units of squared logged income, which exaggerates year-to-year

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13 This trend line and subsequent ones include a number of improvements on the “best replication” figures that have small effects on the results (e.g., better inflation adjustment, and application of sample weights). I settled on a bottom code of $4,000 after comparing the effect on the trend of using bottom codes of $500, $1,000, $2,000, …, $9,000 and determining the lowest bottom code that had a meaningful effect on the estimated trend. Similarly, I tried trimming the bottom 1%, 2%, …, 9% of the data.

14 Alimony income in the PSID was included in transfer income prior to 1994 but was included in taxable income thereafter. Because Hacker combined taxable income from the PSID to transfer income in the CNEF (which includes alimony), alimony is double-counted from 1994 onward. See Lillard (2008) and the documentation in the online PSID Data Center.

15 Since the pre-tax family income variable is from the publicly-available PSID data, I can extend the trend back to 1974. For these estimates, I use a 4-year lag in computing covariances, rather than the 5-year lag in the estimates of mine intended to be consistent with Hacker’s.
differences. When the estimates are presented as standard deviations rather than variances, as shown in Figure 3.3, the increase in volatility over the period is 28 percent over the 1974-2002 period, rather than the 200 percent increase implied by Hacker’s GRS figures.

The next iteration of results from Hacker arrived in the “Revised and Expanded Edition” of Great Risk Shift, published in early 2008.16 This time, rather than volatility increasing over 360 percent between the early 1970s and its 1990s peak, Hacker showed a rise one-third that (less than 125 percent). The new pre-1990s trend was noisier than Hacker’s original estimates, with a sizable decline in volatility during most of the 1980s after the run-up in the first years of the decade. The trend from the early 1990s onward showed a large increase in volatility followed by a large decline in the mid-1990s, with another notable increase from 1998 to 2002. From 1973 to 2004, Hacker showed an increase in volatility of about 95 percent (presented as variances), consistent with my corrected estimates from Figure 3.3.

Once again, Hacker changed the presentation of his data, this time using a chart that treated the 1973 volatility level as a baseline and showing the percent increase in volatility since 1973 on the vertical axis. Figure 3.4 carries over my nationally-representative-subsample trend (measured as variances—the bottom line of the figure).

16 Hacker (2008). Some information on the methods is available at http://pantheon.yale.edu/~jhacker/method.html. Hacker switched to the pre-tax family income variable in the PSID and did not rely on the CNEF at all in his revised analyses. He again adjusted income for family size by dividing by the square root of family size, and he then logged the measures and adjusted them for inflation (using the CPI-U this time). His sample was confined to adults 25-61, including the entire core sample. This time he applied the sample weights. He dropped all non-positive incomes, then he trimmed the top and bottom 1% of the remaining observations. Finally, in computing covariance terms, Hacker used 4-year lags.
from Figure 3.3, presenting it in two different scales.\textsuperscript{17} In the first scale, I show the trend in units of squared logged income on the left axis, as Hacker did in the first edition of \textit{GRS}. The second scale, on the right axis, is the same as that in the revised edition of \textit{GRS}. Both scales extend from 0 to (roughly) the maximum volatility level observed. My estimates resemble the broad Hacker trend, though there are year-to-year differences. The point of Figure 3.4 is to show that Hacker’s presentation change has the effect of visually exaggerating the increase in volatility over time. The adjusted scale makes the vertical distance between 1973 and 2004 volatility appear larger than it does using the unadjusted scale.\textsuperscript{18}

\textbf{Figure 3.4. Great Risk Shift -- Replication of Revised Estimates}

\begin{figure}[h]
\centering
\includegraphics[width=\textwidth]{great_risk_shift}
\end{figure}

\begin{itemize}
\item Footnote 1 in the introduction to the revised edition of \textit{GRS}, as well as footnote 27 of Chapter 1, indicates that all code used to produce the estimates will be publicly available on the book’s website (www.greatriskshift.com), but as of this writing, no code has appeared.
\item The fundamental problem with the scale he uses is that it obscures the possibility of volatility falling below its initial level.
\end{itemize}
Just months after the revised edition of GRS, Hacker and his new collaborator, Elisabeth Jacobs, produced yet another set of estimates, for an Economic Policy Institute research brief. In this most recent version of Hacker’s figures, volatility is shown rising 150 percent from 1973 to 1993, falling by more than half between 1993 and 1998, then increasing between 2000 and 2002. Over the entire 1973-2004 period, volatility essentially doubled. The early-1980s increase in volatility is not as sharp as in the revised edition of GRS, but the latest figures still imply that volatility increased by something like 75 percent just between 1991 and 1993. Hacker and Jacobs acknowledge that “expressed in terms of standard deviations, the 99% rise in transitory variance found from 1973 to 2004 would be around 40%—very close to what is found using [an alternative approach].” Nonetheless, they focus on the variance estimates, showing trends by education group and trends in men’s earnings volatility in terms of variances. The exception—again convenient to Hacker’s argument—is that when they compare overall income dispersion in the PSID to income dispersion in the CPS, they use standard deviations, which reduces differences between the two datasets during the early 1990s. Presenting variances would make the disparity between the CPS and the PSID

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19 Hacker and Jacobs (2008). Once again, Hacker’s analyses rely on the PSID’s pre-tax family income variable. Incomes are again adjusted for family size by dividing by the square root of family size, and they are once again logged and adjusted for inflation using the CPI-U. Again he looks at individuals age 25-61, including the entire core sample and using the sample weights. He dropped all incomes of $1 or less, then trimmed the top and bottom 1% of the remaining observations. Once again, four-year lags are used in computing covariance terms.

20 I have requested the full code used to produce the results in Hacker and Jacobs (2008), as footnote 15 of their brief indicates it is available on request. Unfortunately, I have yet to receive a response after several months.

21 Hacker and Jacobs (2008), page 8. This point is also acknowledged in a long footnote to Chapter 1 of the revised edition of GRS.
look larger and reduce confidence in the large increase in volatility they find during that period.22

All of these methodological problems and presentation decisions aside, there is a broader reason that Hacker’s estimates relying on the Gottschalk-Moffitt model may be problematic, namely that there are good theoretical reasons for believing the model is inadequate for measuring the volatility of household- or family-level income. I return to this question below.

Other Research on Income Instability Trends

As in the previous chapter, I summarize my review of the literature on income instability trends here and provide a longer review in Appendix Two.

Short-Term Directional Mobility

A number of studies have looked at either relative or absolute family income mobility, upward or downward, over short periods. Gottschalk and Danziger (1998) found declines in downward relative mobility from the top quintile in the 1970s and 1980s, and Carroll et al. (2007) confirmed the 1980s decline. Slemrod (1992), however, generally found increasing downward mobility from the top decile, ventile, and centile from the early 1970s through the mid 1980s. Carroll et al. found no change in downward mobility in the early 1990s. Finally, upward relative mobility from the bottom quintile

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22 For a preview of this CPS-PSID comparison, see Figure 3.5 below. A final deceptive presentation decision made by Hacker and Jacobs is to give the same horizontal distance in their charts to two-year differences at the end of the time series as they give to one-year differences over the rest of the time series, which compresses the horizontal space in the charts relative to the vertical space, thereby making volatility increases seem steeper.
was either unchanged or declined in the 1970s and 1980s, and in the first half of the 1980s in particular.\textsuperscript{23} Carroll et al. again found no change in the early 1990s.

There is little consistency across studies in estimated trends in absolute income mobility. The one point on which the research agrees is that upward and downward absolute mobility increased in the 1970s.\textsuperscript{24} Estimates from SSA-based data consistently show only small changes in mobility, generally in contrast to the results from the PSID-based studies.\textsuperscript{25} But the PSID-based studies often disagree among themselves. Most of the evidence from the PSID indicates an increase in both upward and downward absolute mobility over the 1980s and an increase in downward mobility during the late 1980s. The PSID-based studies all agree that mobility was higher in 2000 than in 1970, and several find clear counter-cyclical patterns in directional mobility, in contrast to the findings for relative directional income mobility.\textsuperscript{26} Little else can be said comparing the relative and absolute income mobility results.

\textit{Short-Term Non-Directional Mobility}

Non-directional relative mobility either declined or was stable in the 1970s and declined in the 1980s.\textsuperscript{27} Carroll and his colleagues found an increase in mobility over the first half of the 1990s, though mobility remained below its 1980 level. Non-directional

\begin{footnotesize}
\begin{enumerate}
  \item Gottschalk and Danziger (1998); Gottschalk, McLanahan, and Sandefur (1994); Carroll et al. (2007); Duncan, Smeeding, and Rodgers (1993).
  \item Hacker (2008); Jacobs (2007); Hacker and Jacobs (2008); Jacobs (2008); Gosselin (2008); Gosselin and Zimmerman (2008); Dynan et al. (2008).
  \item Dahl et al. (2008); Orszag (2008).
  \item Hacker (2008); Jacobs (2007); Hacker and Jacobs (2008); Jacobs (2008); Dynan et al. (2008).
  \item Gottschalk and Danziger (1998); Gittleman and Joyce (1999); Carroll et al. (2007).
\end{enumerate}
\end{footnotesize}

I found just one study that examined trends in short-term intertemporal income associations. Tom Hertz examined household income mobility using matched CPS files.\(^{28}\) He found that mobility (measured by one minus the correlation over adjacent years) increased from 1991 to 1998. It then declined slightly in 2004, though the change was not statistically significant. The 1990s trend is the opposite of what Bhashkar Mazumder (2001) found for men’s earnings and what Kopczuk et al. (2009) found for commerce and industry workers’ earnings, but these are the only comparable studies between the earnings and income literatures using correlations.

**Dispersion of Income Changes**

The few studies on dispersion in income changes agree that income volatility increased in the 1970s and changed little in the 1980s.\(^{29}\) The research based on the PSID finds increases in volatility in the 1990s, but Dahl et al.’s declining trend from the mid-1980s to the early 2000s using Social Security Administration data linked to the SIPP is inconsistent with the PSID results.\(^{30}\) Finally, the trend in volatility in the early 2000s is inconsistent across studies. These general findings are consistent with the analogous literatures on trends in absolute income mobility and trends in male earnings dispersion.

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\(^{28}\) Hertz (2006). The data is from the March supplement to the CPS, using three pairs of years. The matched subsamples are re-weighted to be representative of the first year in each pair. Incomes are adjusted for inflation using the CPI-U-RS.

\(^{29}\) Dynan et al. (2008); Orszag (2008); Hacker and Jacobs (2008); Jacobs (2008); Nichols and Zimmerman (2008).

\(^{30}\) Ibid.; Gundersen and Ziliak (2003); Blundell et al. (2008).
in that all find increases in volatility in the 1970s and all find higher volatility in the early 2000s than in 1970 or 1980 in the PSID. On the other hand, the Dahl et al. research implies declining volatility for both earnings and family income since the mid-1980s.

**Within-Person Income Dispersion**

Research on trends in within-person family income dispersion shows a secular rise in mean volatility from the early 1970s to the early 2000s, with the increase (expressed as variances) ranging from roughly two-thirds to 100 percent. Median volatility increased less dramatically according to Nichols and Zimmerman (2008), but Gosselin and Zimmerman (2008) report a sizeable increase. While not entirely consistent, the studies generally find volatility increasing in the 1970s, 1980s, and 1990s, but flat or declining in the early 2000s. This rise is much clearer than the probable increase in within-person earnings dispersion among men found in the previous chapter, and it is more consistent across time than the increase in income-change dispersion.

**Across-Person Dispersion of Earnings Shocks**

The existing studies in this area all use the PSID and are generally consistent. Volatility increased in the 1970s (except in Gottschalk and Moffitt, 2007), in the 1980s, in the 1990s (except in Nichols and Zimmerman, 2008), and from 1998 to 2002. These results are consistent with the trends in within-person income dispersion and in the dispersion of earnings shocks noted in the previous chapter. The increase from the early

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1970s to the early 2000s was somewhere between 20 and 50 percent expressed as standard deviations, according to these studies. That is comparable to the range found in the literature on male earnings volatility that use the same methods, summarized in Appendix One as running from 15 to 65 percent.

Summary of Previous Literature

The research on family income instability and volatility trends is much more consistent than that on trends for earnings. This fact is largely due to the dominance of PSID-based studies in the research on family income instability and volatility trends, and the relatively small number of studies thus far conducted. The research generally finds that short-term relative income mobility was largely unchanged or declined during the 1970s and the 1980s. However, research on absolute mobility and on income dispersion show that family income instability and volatility increased in the 1970s, 1980s, and 1990s. There is conflicting evidence for the early 2000s, with research on dispersion of transitory shocks showing an increase and that on within-person income dispersion showing a flat or declining trend.

The big exception to these conclusions is the CBO research, which uses the SIPP linked to Social Security records to correct for measurement error in labor income. The CBO research indicates that income instability was basically flat from the mid-1980s forward. Otherwise, the increases in income volatility since 1970, 1980, or 1990 are consistent with the increases in male earnings volatility over these same periods in the research discussed in Appendix One. Where quantitative estimates of volatility can be
compared between the income and male earnings literatures (studies using dispersion-based measures) the two show increases of similar magnitudes.

In short, while the rise in income volatility has not been nearly as large as Hacker’s initial estimates implied, the evidence from previous studies using the PSID does support his Great Risk Shift hypothesis in that it generally finds steadily increasing income volatility and instability over the past thirty-five years. Some dispersion-based estimates imply a steady rise in volatility of as much as 50 percent over the entire period, although the CBO figures suggest that it did not change at all after the early 1980s. This range is uniformly much lower than the doubling of volatility that Hacker has reported, but it implies potentially high volatility nonetheless. The rest of this chapter attempts to make more sense of the size and timing of any rise in income instability.

**Methods and Data**

To determine whether income volatility and instability trends are consistent using a variety of measures, a single dataset, and a standard set of carefully thought-through methodological decisions, and to see whether these trends are similar to volatility trends for earnings, I estimated several trends using the Panel Study of Income Dynamics (PSID). To the extent possible, I made the same methodological decisions as in Chapter Two. I also build gradually from the earnings volatility estimates by beginning with aggregated family earnings volatility trends and working up toward alternative family income concepts. Below I briefly review the key methodological points for understanding the estimates that follow. See the Methods and Data section of Chapter Two for a full discussion of the considerations involved.
Data

As in Chapter Two, I use only the nationally-representative “SRC sample” of the PSID, excluding both the disadvantaged SEO sample and the Latino and immigrant samples of the 1990s. The unweighted SRC sample consistently produced income variance trends and levels that were closer than estimates from the weighted “core” PSID sample (SRC plus SEO) to the Annual Social and Economic Supplement to the Current Population Survey. Figure 3.5 shows the SRC vs. CPS comparisons. I also ran my variance decomposition models using the weighted core sample, and the differences were not large.

One important change that I make in this chapter is to include “non-sample” members if they are in a family unit that has at least one sample member and they otherwise meet the sample restrictions. I do so because my unit of analysis consists of persons rather than families. If non-sample members are excluded, that will disproportionately remove partners of sample members in more recent years of the PSID. In the 1968 wave, both partners in a sample family unit are defined as sample members. But in subsequent years, people with no blood ties to an earlier sample member are not designated as sample members. In particular, when the children of the first PSID generation marry or cohabit with someone outside the PSID sample, the partner is deemed non-sample.

33 I obtained the CPS data from Unicon Research Corporation (www.unicon.com). I compared the PSID variances of logged family income for heads of family units and logged household income for heads of household units (age 20-59) with CPS variances of family income for heads of families and household income for heads of households (age 20-59). I also used different trims of the top and bottom of these income distributions. Figure 3.5 presents results trimming the top and bottom 2 percent of incomes. Full results available from the author upon request.
If families were my unit of analysis, this would not be a problem because the ratio of families with partners to single-person family units would be unaffected by the exclusion of non-sample partners. But since persons are the unit of analysis, excluding non-sample members will reduce the ratio of adults in families with partners to adults without partners. Note that non-sample members are assigned weights of zero, so that analysts who simply use the PSID sample weights exclude these individuals automatically. This probably explains some of the differences in results discussed in Appendix Two, since excluding these non-sample members makes trends look worse than they were by overstating the fraction of adults without partners. Including non-
sample adults in family units with a sample member produces variance trends that more closely match CPS variance trends (results not shown).

Chapter Two details the changing definition and coding of several earnings variables. For this chapter, I created three different measures of combined head and spouse earnings, each of which allocates income from self-employment into labor and asset components differently to deal with changes in the way the PSID coded these variables.

Chapter Two also discusses changes in the PSID since 1990. These changes include:

- The switch to more detailed questions for certain income components in 1993;
- Updated data collection, editing, and processing procedures in 1993, 1994, 1995, and 2003;
- Cuts to the editing budget beginning in 1993;
- Doubling of the interview length between 1995 and 1999;
- Changes in the rules for following family members who move out, which occur in 1990, 1993, 1994, 1996, 1997, and 2005; and
- A change in the sample definition to include more children in 1994.

These changes appear to have affected the PSID estimates. Research by the Institute for Social Research has shown that the introduction of computer-assisted telephone interviewing (CATI) and income processing software in 1993 increased income variances, and income variance in the 1993 PSID survey is not only greater than in the 1992 or 1994 PSID, it is suspiciously large compared with the 1993 CPS.
Percentiles below the median show one-time declines, and those above the mean show increases.\textsuperscript{34} The PSID income estimates at the first, third, and fifth percentiles are below those from the CPS from 1992 to 1996, even though the opposite is true in all other years between 1967 and 2004 and for all but the lowest percentiles even between 1992 and 1996.\textsuperscript{35} At the very bottom of the income distribution (below the tenth percentile), the PSID and CPS income trends do not match up very well between the mid-1980s and the early 1990s, with the PSID percentiles declining anomalously. There is an up-tick in income for these percentiles in the 2004 PSID but not in the CPS. Hacker and Jacobs (2008) show an increase in the dispersion of income in the PSID in the early 1990s that is not reflected in the CPS. So does my Figure 3.5.\textsuperscript{36} Dynan et al. (2008) discovered a sizable jump beginning in the early 1990s in the frequency of heads who report working more than 120 hours but report $0 in earnings. Nichols and Zimmerman (2008) report a similar jump in the early 1990s.

I take several precautions that help mitigate the effects of these changes. First, I trim the top and bottom 2 percent of positive family incomes in each year, which reduces the impact of changes in bottom codes.\textsuperscript{37} I minimize the influence of very high incomes by using the natural log transformation in most analyses, making my results less sensitive to large incomes and large changes in high incomes. I address the changes in recontact

\textsuperscript{34} Kim and Stafford (2000), Gouskova and Schoeni (2007).
\textsuperscript{35} Gouskova and Schoeni (2007).
\textsuperscript{36} See also http://pantheon.yale.edu/~jhacker/cps_psid.pdf for a version of Hacker and Jacob’s chart with variances compared.
\textsuperscript{37} That is, the top and bottom 2 percent of families – rather than individuals – are trimmed in each year. All individuals in families not trimmed are assigned the income of their family. This trimming occurs within age groups defined by the head’s age, with the categories consisting of heads age 21-30 years, 31-40 years, 41-50 years, and 51-60 years. I relied on my comparisons with CPS income variances to determine what trims to use.
rules and sample definitions, by re-running my variance decompositions restricting the sample to exclude the added respondents. The results were not meaningfully affected. Nor were the results affected when I excluded individuals with incomes that included imputed components.

One way to assess how important these changes to the PSID are is to compare trends for income measures that aggregate fewer and greater numbers of variables. Measures that aggregate a relatively large number of variables are more likely to be affected by data inconsistencies. The measures I examine below provide a range of aggregation to consider this issue.

I adjust all incomes for inflation using the CPI-U-RS, linking it to the CPI-U for earlier years. In most analyses I transform income by taking its natural log. Using natural logs has the advantage of transforming income to a scale that is mean-independent. An across-the-board 10 percent change in logged income will not produce a change in the logged income variance. But a problem with the log transformation is that it has the effect of increasing the influence of very small incomes and very small absolute changes in low incomes, as discussed in regard to Jacob Hacker’s analyses above. Because a non-negligible number of PSID respondents report very low family incomes, year-to-year changes in measured volatility can be strongly influenced by changes in the proportion of individuals reporting very low incomes. In the PSID, there is enough year-to-year variation in low reported incomes to drive estimated changes in volatility over time. Trimming very low incomes before estimating volatility reduces this problem.

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38 See [http://www.census.gov/hhes/www/income/income07/AA-CPI-U-RS.pdf](http://www.census.gov/hhes/www/income/income07/AA-CPI-U-RS.pdf). I also ran variance decomposition models using the CPI-U and ran all models using the CPI-U-RS linked to the CPI-U-X1, and the results were similar.
My analyses include only persons with positive income in the years considered. This restriction is partly for methodological reasons – the logs of zero and negative numbers are undefined. But persons who report going an entire year without any family income are also likely either to have special circumstances that make their inclusion in these analyses inappropriate or to be under-reporting their income. Chapter Two shows that just one in twenty people with no earnings in an entire calendar year report that they want a job and are having trouble finding work. Since several of my income measures include public transfers, it is likely that a large proportion of $0 income reports are simply wrong. However, individuals with business losses can and do report zero or negative income. As a fraction of the sample, though, adults with asset income losses who report low incomes are a tiny group.

I attempt to minimize the number of retirees and students by restricting my sample to persons between the ages of 21 and 60 in the survey year (which makes most of them 20 to 59 years old in the year for which income is reported). I chose this range as a compromise between being more inclusive and keeping sample sizes large, and wanting to exclude students and retirees. Since labor force status is reported for the current week while income is reported for the previous year, identifying individuals who were students and retirees during the income year is impossible after 1997, when the PSID became biennial. Furthermore, wives’ labor force status is unavailable in the PSID before 1976.

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39 Edin and Lein (1997). Including persons without income in the variance decomposition models is problematic, even if their incomes are recoded to $1 or some other value. The model requires permanent and transitory components to sum to total income. A person with $0 in income must have both $0 in permanent income and $0 in transitory income or must have a transitory component that exactly cancels out his or her permanent component. The distribution of transitory income, in expectation, should be the same among those who do not work as it is among those who do. When one attempts to estimate the variance decomposition model by including those with no income (after recoding their incomes to be positive), the transitory variance estimates balloon. There are similar theoretical problems with recoding $0 incomes when using error components models.
(and also in 1977 and 1978). However, I ran the variance decomposition models excluding people who were students or retired in both the current survey and two years earlier, bracketing the year for which income is reported. I also ran all variance decomposition models restricting the sample to individuals 18 to 64 years old. The results were negligibly affected.40

The PSID asks about current family unit members’ income during the previous year. If family composition has changed, the sum of these amounts may not equal the total income for the previous year’s family unit members. In particular, if a couple divorces or gets married, there will be a mismatch between the previous and the current year’s family members. The two groups will be the same only for those families that neither gained nor lost earners. As a result family incomes are usually mis-measured when adult family composition changes.

In order to test the sensitivity of my results to this issue, I repeated selected analyses for a restricted sample confined to adults in families where the head was married or single and there was no marriage or divorce in either the previous or current calendar year.41 This restriction is not ideal. Limiting family composition change removes an important source of volatility. Furthermore, the sample restriction introduces selection into the results in two different ways. First, family composition may be affected by

40 Note that while I exclude persons under age 20 or older than age 59, the income measures that I use that include income from all family members include income from people under age 20 or older than age 59. Additional sample restrictions used include requiring individuals to be present in a family in the survey year (not necessarily the same family when comparing multiple years) and requiring them to be PSID sample members.

41 I chose a relatively conservative restriction given that family composition can change in the previous calendar year either before or after PSID respondents report the composition of their family at the time of the interview and given that it can change in the current calendar year either before or after the PSID interview. This restriction does not take into consideration the possibility that cohabiters move in or out while the head’s marital status remains single. Nor does it consider changes in the number of people receiving income other than the head or partner.
income volatility rather than the other way around. Second, measures of family structure and marital history based on calendar years—and collected retrospectively for the “off years” even after the PSID became biennial—are obtainable only using the supplemental PSID Family History files, which did not assess family and marital histories until 1985. Respondents lost before 1985 never provided their histories, so they are excluded. These qualifications aside, the results of these sensitivity analyses can be informative as to whether mismatches related to measuring family composition changes are potentially important. Note that if the mismatch problem has a constant effect on volatility estimates across all the years in the PSID, the trend will be unaffected (although the percent changes could be). In the analyses below, none of my conclusions are much affected by this additional sample restriction.

**Income Measures**

I examined several income measures, four of which are included in the analyses below. I begin by combining the labor income of heads and partners. This measure includes “labor” income from self-employment, and excludes the share of self-employment income that the PSID labels “asset” income. I then re-estimated my results using a measure that assigned all self-employment income to labor. The results were unaffected.\(^4^2\)

I also show results for pre-tax family income and post-tax family income. Pre-tax income in the PSID is a composite variable that includes all labor, asset, and transfer

\(^4^2\) I also examined two other versions of the measure. One excluded heads’ non-wage labor income. Another dropped the labor part of income from taking in roomers and boarders and added the part of gardening income assigned to asset income, in order to take into account changes in the PSID definition of labor income over time. Neither alteration affected the results.
income of all family unit members age 16 or older living in the household. It does not include capital gains, lump sum payments such as lottery winnings, or in-kind transfers. However, I also computed a measure that included lump-sum income and one that included the value of food stamps, public energy subsidies, and free housing. The results were unaffected. I omit them here because these additional variables are available in fewer years.43

To estimate post-tax income, I relied on an enhanced version of the PSID created by Cornell University’s Department of Policy Analysis and Management, the Cross-National Equivalent File (CNEF).44 The CNEF includes estimates of federal and state income and payroll taxes, based on the National Bureau of Economic Research’s TAXSIM model.45 I also created a comprehensive post-tax-and-transfer measure that included food stamps, energy subsidies, and the value of free rent. Once again, the results were unaffected.46

Finally, I include estimates using pre-tax income adjusted for family size. To adjust incomes, I divide them by the square root of family size (prior to logging income for those analyses in which logged income is used). Reported family size in the PSID

43 I recoded all incomes of $1 or less to $0 to account for coding changes in the PSID (incomes were bottom-coded at $0 or $1 prior to 1994). In the 1999, 2001, and 2003 surveys, after the PSID switched to biennial surveys, families were asked about income received two years prior to the survey. I attempted using these income measures to fill in the gaps created by the switch to biennial surveys, but the results clearly showed that the variables were not comparable to the one-year-recall questions. I also tried a version of pre-tax income that simply added heads’ and partners’ combined taxable income, other family members’ taxable income, heads’ and partners’ combined transfer income, and others’ transfer income. The results were hardly affected.


45 Butrica and Burkhauser (1997).

46 While I omit the results for brevity, I also examined volatility trends in heads’ and partners’ combined taxable income (including labor and asset income) and total family taxable income. The trends mirrored those shown for the other income measures below. Results available from the author on request.
refers to the current year’s family, while reported income is that received by current family members in the previous year. Since the adjustment of last year’s income should be based on last year’s average family size, changes in marital status, births, deaths, and changes due to family members moving out (or back in) all introduce error into the family size adjustment. Nearly all other annual household surveys—as well as the decennial census—suffer from this problem.

To check the sensitivity of my results to these issues, I estimated trends in adjusted pre-tax income using a second approach—dividing income by needs, based on food budgets prepared by the U.S. Department of Agriculture. The benefit of this measure is that it accounts for family composition in the previous year (the same year income is measured) rather than simply reflecting the current year’s family size.47 The downside is that for some measures of income instability, using a ratio of income to needs rather than some version of per-capita income may be inappropriate. Dividing by the square root of family size leaves the units in dollars, but dividing by income needs produces a unit-less ratio.

I conducted a second sensitivity test to see whether my results hold when I adjust by the square root of family size but confine the analyses to adults in families with continuously married or single heads over the previous and current year who had no births to the head or partner. The results of these tests indicated that none of my conclusions are affected by the different adjustment or sample selection.

47 University of Michigan Survey Research Center (1972), page 302.
Measures of Volatility and Instability

The measures used in this chapter parallel the earnings volatility and instability measures from Chapter Two.

Relative Mobility. I examine the probability of rising into or falling out of a quintile over two years.\textsuperscript{48} I compute quintiles from positive trimmed income for each sample and each income measure, in every year. That is, quintiles are constructed after excluding persons outside the sample of interest (e.g., persons under age 20, persons with no income), but without regard to whether a person’s income is observed in any other year (e.g., the second year over which mobility is measured when constructing quintiles for the first year). This approach means that quintile definitions are not fixed across analyses. It also means that upward mobility from a quintile need not be exactly matched by someone moving downward into it.

Absolute Mobility. I compute two measures of absolute mobility: the probability of experiencing a loss or gain of 25 percent or more over two years, and the probability of experiencing a loss or gain of $10,000 or more (in 2007 dollars). I do not log income for these analyses. Because I exclude those with non-positive income, I avoid the methodological problem of how to calculate the percentage increase when income rises from zero to a positive amount. Because I trim the bottom and top two percent of observations, I also avoid treating small increases from very low initial incomes as large percent changes (e.g., treating a change in income from $100 to $200 as equivalent to a change from $100,000 to $200,000).

\textsuperscript{48} A window of two years is the smallest that can be considered over the entire course of the PSID because the survey switched to biennial interviews after 1997.
Non-directional Mobility. I proceed from the directional mobility results to trends in the likelihood of moving either up or down (in terms of quintiles or in terms of a 25 percent change in absolute income). I also include two measures of intertemporal income association: the Pearson product-moment correlation coefficient of logged income separated by two years, and Spearman’s rank correlation coefficient.

Dispersion of Income Changes. Following several previous studies, I compute the standard deviation of two-year changes in logged income. This operationalization of instability essentially extends the absolute mobility analysis by looking at the full distribution of income changes. I use the difference in logged incomes rather than looking at percent changes because of the asymmetry involved in converting income increases and drops to percentages. Imagine a group of three workers making $20,000, $30,000, and $40,000 in one year who each see their incomes rise by $10,000 the next. The percent changes are 50%, 33%, and 25%. The next year their earnings each fall by $10,000 – constituting percent changes of -33%, -25%, and -20%. The standard deviation of percent changes has declined, even though incomes simply returned to their original levels.\(^{49}\) This problem does not arise with changes in log income.

Within-Person Income Dispersion. For every sample member, I estimate the standard deviation of their logged income across five years in a nine-year window. I then treat the mean of these individual standard deviations as a summary of volatility in the year on which the nine-year window is centered. Specifically, I follow Gosselin and Zimmerman (2008) in using income in years \(t-4, t-2, t, t+2,\) and \(t+4\) and report volatility in year \(t\) as the average standard deviation of income over these five years. The incomes

\(^{49}\) Dynan et al. (2008) and Dahl et al. (2008) address this problem by using the average of the two years as the denominator in computing percent changes, though Dahl et al. also present results in an appendix using my approach.
are separated by two years because the PSID switched to biennial surveys after 1997. Like Gosselin and Zimmerman, I require a person to have at least three years of income out of the five years.

**Dispersion of Transitory Income Shocks.** As discussed in Chapter Two, I include two different model-driven measures of the dispersion of transitory income shocks. These measures depend on assumptions about income dynamics that hold imperfectly at best.

My first set of transitory income dispersion measures is based on the variance decomposition model of Peter Gottschalk and Robert Moffitt (see the discussions in Chapter Two and Appendix Two).\(^5\) The variance decomposition model—which Gottschalk and Moffitt now reject but which has remained fairly popular—relies on a simple model of earnings dynamics that allows the total variance in earnings to be expressed in terms of permanent and transitory variance. The model essentially specifies earnings to be static in the short-run, except for random annual shocks that completely dissipate over the short-run. Volatility is operationalized as dispersion of the shocks.

When applied to *family income* dynamics, the model’s assumptions are even less realistic than they are for earnings. As discussed in Chapter Two, the Gottschalk-Moffitt model of individual earnings dynamics begins with a model of earnings in dollars, not yet logged:

\[
\pi_t = \pi_t^e \omega_t, \tag{1}
\]

---

Total earnings $z_i$ are expressed as the product of a permanent component $\pi_i$ and a transitory component $\omega_i$ that is a random shock to permanent earnings. Taking the natural log of both sides of Equation (1) leads to the transformed equation:

$$\log(z_i) = \log(\pi_i \omega_i) = \log(\pi_i) + \log(\omega_i),$$

which can be rewritten as

$$y_i = \mu_i + v_i.$$

To estimate the variance of the transitory component, one takes advantage of the fact that if the $v_i$'s are random shocks, then the overall earnings variance is

$$Var(y_i) = Var(\mu_i + v_i) = Var(\mu_i) + Var(v_i).$$

Furthermore, the covariance of overall earnings measured $k$ years apart is

$$Cov(y_i, y_{i-k}) = Cov(\mu_i, v_i, \mu_i, v_{i-k}) = Cov(\mu_i, v_i) + Cov(\mu_i, v_{i-k}) + Cov(v_i, v_{i-k}) = Var(\mu_i).$$

Because one can estimate $Var(y_i)$ and $Cov(y_i, y_{i-k})$, one can estimate the transitory variance $Var(v_i)$.

Now consider what happens when one attempts to apply the model to family-level economic volatility analyses. For simplicity, assume that all income consists of earnings, and that all families consist of male-female couples and no other earners. Now $i$ indexes couples rather than individuals, and $h$ and $w$ denote the husband and wife, respectively. The model in Equation (1) now becomes

$$z_h^h + z_w^w = \pi_h^h \omega_h^h + \pi_w^w \omega_w^w.$$

But now the log transformation gives

$$\log(z_h^h + z_w^w) = \log(\pi_h^h \omega_h^h + \pi_w^w \omega_w^w) \neq \log(\pi_h^h + \pi_w^w) + \log(\omega_h^h + \omega_w^w).$$
Equation (2b) does not allow us to use a neat model in which permanent and transitory components are separate and additive. What does one end up with if one subtracts $\text{Cov}(y_{it}, y_{it-k})$ from $\text{Var}(y_{it})$? It is straightforward to see if we express wives’ permanent earnings in terms of their husbands’, so that

$$\pi_i^w = \theta_i \pi_i^h.$$  

(6)

Equation (1b) can now be rewritten as

$$z_i^h + z_i^w = \pi_i^h \omega_{it}^h + \theta_i \pi_i^h \omega_{it}^w = \pi_i^h (\omega_{it}^h + \theta_i \omega_{it}^w),$$  

(1c)

and now we can take the natural log of both sides to get $y_{it} = \mu_i + \nu_{it}$, where all three variables are logs, with

$$\mu_i = \log(\pi_i^h),$$ and

$$\nu_{it} = \log(\omega_{it}^h + \theta_i \omega_{it}^w) = \log(\omega_{it}^h + (\pi_i^w / \pi_i^h) \omega_{it}^w).$$  

(8)

But now $\nu_{it}$ and $\mu_i$ may be correlated and the $\nu_{it}$’s may also be due to the $\theta_i$ term. The variance of $y_{it}$ becomes

$$\text{Var}(y_{it}) = \text{Var}(\mu_i + \nu_{it}) =$$

$$\text{Var}(\log(\pi_i^h)) + \text{Var}(\log(\omega_{it}^h + (\pi_i^w / \pi_i^h) \omega_{it}^w)) + 2 \text{Cov}(\log(\pi_i^h), \log(\omega_{it}^h + (\pi_i^w / \pi_i^h) \omega_{it}^w)).$$  

(4c)

The covariance term becomes

$$\text{Cov}(y_{it}, y_{it-k}) = \text{Var}(\log(\pi_i^h)) + \text{Cov}(\log(\pi_i^h), \log(\omega_{it}^h + (\pi_i^w / \pi_i^h) \omega_{it}^w)) +$$

$$\text{Cov}(\log(\pi_i^h), \log(\omega_{it-k}^h + (\pi_i^w / \pi_i^h) \omega_{it-k}^w)) +$$

$$\text{Cov}(\log(\omega_{it}^h + (\pi_i^w / \pi_i^h) \omega_{it}^w), \log(\omega_{it-k}^h + (\pi_i^w / \pi_i^h) \omega_{it-k}^w))$$

(5c)

$$\neq \text{Var}(\log(\pi_i^h + \pi_i^w)).$$

And the difference between the variance and covariance terms is
Equation (9) implies that the estimated transitory variance of husbands’ and wives’ logged combined income depends on the joint distribution of the ratio of wives’ to husbands’ permanent income (assortative mating) and of husbands’ and wives’ transitory logged earnings. The direction of bias is indeterminate.

Only if both husbands’ and wives’ earnings are subject to identical shocks in every year will the variance decomposition model for earnings apply to family-level earnings. In that case, Equation (1b) becomes

\[ z_{it}^h + z_{it}^w = \pi_t^h \omega_t^h + \pi_t^w \omega_t^w = \omega_t (\pi_t^h + \pi_t^w) \]  

(1d)

and Equation (2b) becomes

\[ \log(z_{it}^h + z_{it}^w) = \log(\omega_t (\pi_t^h + \pi_t^w)) = \log(\pi_t^h) + \log(\pi_t^w) + \log(\omega_t). \]  

(2d)

An alternative to formulating the variance decomposition model in terms of shocks to earnings is to formulate it in terms of shocks to family income, in which case Equations (1d) and (2d) also apply. But then one must assume that all income components are static—the earnings of both partners, transfers, and capital income—or continually adjusted to be static in the aggregate within the window from which covariances are estimated, except that the family is subject to annual random shocks that do not persist from year to year. That rules out any coordination between husbands and
wives in terms of hours or the type of work each does. For example, temporary
departures from the labor force—say, to raise children—are ruled out.

What if the Gottschalk-Moffitt model does not describe the actual income
dynamics of families? As noted in Chapter Two, Shin and Solon (2009) and Moffitt and
Gottschalk themselves have argued that the model may be too simple to be useful even in
the context of individual earnings dynamics.\textsuperscript{51} If the true model of earnings dynamics is
only slightly more complicated than the Gottschalk-Moffitt model, then estimates based
on the model may be substantially biased—and the bias is likely to be even larger when
the model is applied to family-level income.

Shin and Solon outline an alternative model that is only slightly more complex
than Gottschalk and Moffitt’s and show that it implies their model may be inadequate.
Assume that the analyst proceeds to estimate earnings volatility as if Equation (1) were
the true model but that the actual true model is
\[ z_{it} = (\pi_i)^{\rho_i} \omega_{it}. \] (1e)
Taking the log of both sides yields
\[ \log(z_{it}) = \log((\pi_i)^{\rho_i} \omega_{it}) = \log((\pi_i)^{\rho_i}) + \log(\omega_{it}) = \rho_i \log(\pi_i) + \log(\omega_{it}), \] or (2e)
\[ y_{it} = \rho_i \mu_i + v_{it}. \] (3e)
This time, the individual effect is subject to changing economic returns over time.
Estimating the variance of overall earnings now gives
\[ Var(y_{it}) = Var(\rho_i \mu_i + v_{it}) = \rho_i^2 Var(\mu_i) + Var(v_{it}), \] (4e)
and the estimated covariance is

\[ \text{Cov}(y_{it}, y_{it-k}) = \text{Cov}(\rho_{it}, y_{it} + v_{it}, \rho_{t-k} \mu_{t} + v_{it-k}) = \]
\[ \rho \rho_{t-k} \text{Var} \mu_{t} + \text{Cov}(\rho_{it}, y_{it}, v_{it} + \text{Cov}(\rho_{t-k} \mu_{t}, v_{it}) + \text{Cov}(v_{it}, v_{it-k}) = \]
\[ \rho \rho_{t-k} \text{Var} \mu_{t}. \]

Subtracting Equation (5e) from Equation (4e) results in an estimate of transitory variance equal to
\[ \text{Var}(v_{it}) = \text{Var}(v_{it}) + \rho_{t} (\rho_{t} - \rho_{t-k}) \text{Var} \mu_{t}. \] (10)

The estimate will be biased by the second term unless \( \rho_{t} = \rho_{t-k} \). And if volatility estimates from two adjacent years are compared, the bias in the change estimate is
\[ [\rho_{t} (\rho_{t} - \rho_{t-k}) - \rho_{t-1} (\rho_{t-1} - \rho_{t-1-k})] \text{Var} \mu_{t}. \]

One can repeat this exercise for family-level volatility estimates. The true model is
\[ z_{it}^{h} + z_{it}^{w} = (\pi_{it}^{h})^{\phi} \omega_{it}^{h} + (\pi_{it}^{w})^{\phi} \omega_{it}^{w} = (\pi_{it}^{h})^{\phi} \omega_{it}^{h} + \theta_{it} (\pi_{it}^{h})^{\phi} \omega_{it}^{w}, \]
\[ = (\pi_{it}^{h})^{\phi} (\omega_{it}^{h} + \theta_{it} \omega_{it}^{w}) \] (1f)

where this time
\[ \theta_{it} = (\pi_{it}^{w})^{\phi} / (\pi_{it}^{h})^{\phi} \] (6f)

and is subscripted by time, since the returns to fixed effects are time-varying. Once again, taking the log of both sides yields
\[ \log(z_{it}^{h} + z_{it}^{w}) = \log((\pi_{it}^{h})^{\phi} (\omega_{it}^{h} + \theta_{it} \omega_{it}^{w})) = \log((\pi_{it}^{h})^{\phi}) + \log(\omega_{it}^{h} + \omega_{it}^{w}) + \]
\[ \rho^{h} \log(\pi_{it}^{h}) + \log(\omega_{it}^{h} + \theta_{it} \omega_{it}^{w}) \] (2f)

which can again be expressed as
\[ y_{it} = \rho_{t} \mu_{t} + v_{it} \] (3f)

where
\[ y_{it} = \log(z_{it}^{h} + z_{it}^{w}) \] (11)
\[ \mu_i = \log(\pi_i^h), \quad (12) \]
\[ \nu_i = \log(\omega_i^h + \theta_i^a \omega_i^w), \text{ and} \quad (13) \]
\[ \rho_i = \rho_i^h. \quad (14) \]

Subtracting the covariance of incomes \( k \) years apart from the variance of income in year \( t \) results in an estimate of transitory variance equal to

\[
\hat{\text{Var}}(\log(\omega_i^h + \omega_{i-k}^w)) = \text{Var}(\rho_i^h \log(\pi_i^h) + \log(\omega_i^h + \theta_i^a \omega_i^w))
- \text{Cov}(\rho_i^h \log(\pi_i^h) + \log(\omega_i^h + \theta_i^a \omega_i^w), \rho_{i-k}^h \log(\pi_{i-k}^h) + \log(\omega_{i-k}^h + \theta_{i-k}^a \omega_{i-k}^w)) =
\text{Var}(\log(\omega_i^h + \theta_i^a \omega_i^w)) + \rho_i^h (\rho_i^h - \rho_{i-k}^h) \text{Var}(\log(\pi_i^h))
+ (2 \rho_i^h - \rho_{i-k}^h) \text{Cov}(\log(\pi_i^h), \log(\omega_i^h + \theta_i^a \omega_i^w)) - \text{Cov}(\rho_i^h \log(\pi_i^h), \log(\omega_{i-k}^h + \theta_{i-k}^a \omega_{i-k}^w))
- \text{Cov}(\log(\omega_i^h + \theta_i^a \omega_i^w), \log(\omega_{i-k}^h + \theta_{i-k}^a \omega_{i-k}^w))
\quad (15) \]

Note that there are now three additional bias terms compared with Equation 10. Note too that the above discussion actually understates the potential for bias in applying the variance decomposition model of earnings dynamics to income dynamics, because it does not even take into account sources of income other than earnings.

Finally, nothing in the above discussion incorporates income adjusted for family size. Thinking about random, non-persistent shocks as applying to family-size-adjusted income makes even less sense than applying the model to non-adjusted family income. The model requires that family income adjusted for size is static within the relevant window, save for the annual shocks. If family size changes in a way that is not simply random, then income must recalibrate so that adjusted income remains constant. If it changes in a way that is random, then the change must reverse itself in the short-term.

Ultimately, whether these modeling issues are important or not is an open question. The point-in-time estimates might all be biased, but unless the bias changes systematically, the trends might still be valid. However, estimated percent changes in
volatility depend on baseline levels as well as the size of change, so such estimates should be viewed cautiously.

As in Chapter Two, I present results using four-year lags to estimate the covariance term in Equation (5). I also adjust incomes to remove the effect of the head’s age by pooling heads across all PSID waves and regressing trimmed logged income on a quartic in age, year indicators, and individual fixed effects (stratified by sex in analyses that combine men and women). I use the age coefficients to age-residualize income for the heads, and I assign this income to all family members who meet the other sample restrictions.

My second measure of transitory dispersion is based on an error components model that is similar to that of Haider (2001). As in Chapter Two, I model income as a function of the head’s age, individual fixed effects, a random growth component (individual-specific slopes), and a time-varying transitory component that follows an ARMA (1,1) process:

\[
y_{it} = f(a_{it}, t) + \mu_i + \gamma t + \nu_{it}
\]

\[
\nu_{it} = \rho \nu_{it-1} + \theta \varepsilon_{it-1} + \varepsilon_{it}
\]  \hspace{1cm} (6)

This model implies that once incomes are age- and year-residualized, which removes the first term on the right-hand side in the equation for \(y_{it}\), the covariance of incomes in any two years \(t\) and \(s\) is given by

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52 I am indebted to Lorenzo Cappellari for sharing his STATA code with me, without which I could not have produced the error components trends below. I require sample members to have at least five non-missing observations in odd-numbered survey years from the time they first appear in the data, and I require them to have no more than 20 percent of observations in odd-numbered survey years missing from the time they first appear. Note that these criteria exclude people who enter the PSID sample in 1999 or later, since they cannot have five non-missing observations.

53 Technically, I regress trimmed logged earnings on a quartic in age and year dummies. The regressions are also stratified on sex for the estimates that combine men and women. As with the variance
\[
\text{Cov}(y_t, y_s) = \text{Var}(\mu_t) + ts\text{Var}(\gamma_t) + (t + s)\text{Cov}(\mu_t, \gamma_t) + \text{Var}(\varepsilon_t) \rho^{t-s}[1 + \theta / \rho + (\rho + \theta)^2 / (1 - \rho^2)] .
\] (7)

This model may be estimated by stacking the elements \( m \) of the empirical covariance matrix (including the diagonal and above), creating variables \( t \) and \( s \) to designate rows and columns, and using nonlinear least squares to estimate the model

\[
m_c = b_0 + tsb_1 + (t + s)b_2 + \{b_3^{t-s}[1 + b_4 / b_3 + (b_3 + b_4)^2 / (1 - b_3^2)]\}(b_5 + \sum_{j=2}^{C} d_{t-s}b_{j+s})
\] (8)

where the vector of dummy variables \( d \) indicate whether \( t = j \). The coefficients \( b_5 \) to \( b_{C+4} \) provide the estimates of each year’s transitory variance, of which I then take the square root to get the transitory standard deviation.\(^{54}\)

**Pivot Volatility.** Once again, I close with trends in what I term “pivot volatility”, which addresses the problem, noted in the introduction, that the existing measures of volatility have in that they may interpret secular increases in income as volatility. As discussed in Chapter Two, the conceptual basis for pivot volatility is that volatility occurs when individuals experience a reversal of their income trajectory. To construct my measure of pivot volatility, I look at the points within a nine-year window where income can reverse its two-year trajectory across five years. Over nine years, there are three such points. At each potential pivot, individuals are assigned the average of their pre- and post-pivot two-year income changes (each expressed as the absolute value of the percentage change) or zero if a pivot did not occur. They are then assigned the average decomposition estimates, this residualization is conducted on family heads, and the income is then assigned to all family members who meet the other sample restrictions.

\(^{54}\) My model differs from that of Haider in that Haider allows the price of fixed effects and slopes to vary over time.
of their three pivot values, and I estimate trends in the mean of this within-person average across individuals. Pivot volatility indicates, roughly, the typical pre- and postpivot percentage change experienced by individuals. For more details, see Chapter Two.55

**Results**

*Short-Term Downward Mobility*

I follow the organization of Chapter Two and first present trends in downward mobility before moving to non-directional measures of instability and volatility. Figure 3.6 is a first step at connecting the earnings results from Chapter Two with the family income results. It shows trends in downward relative and absolute mobility for combined head and partner labor income. Figures 3.7, 3.8, and 3.9 present analogous trends for pre-tax and post-tax family income and for pre-tax income adjusted for family size. Looking first at relative mobility, the probability of combined labor income falling one or more quintiles over two years declined steadily between the late 1960s and the early 2000s. The decline in mobility for the other measures of family income is confined to the late 1970s and 1980s.

These results are broadly consistent with the declines in relative earnings mobility I found for both men and women in Chapter Two. They are also consistent with those of Peter Gottschalk and Sheldon Danziger (1998) in that we both find less downward mobility in the early 1990s than in the late 1960s using family-size-adjusted income, but

---

55 The estimate for year $y$ is technically the mean across people of the average of

$$\text{abs}\left(\frac{y_{t+2}-y_t}{y_{t+2}+y_t}\right) + \text{abs}\left(\frac{y_{t-2}-y_t}{y_{t-2}+y_t}\right),$$

$$\text{abs}\left(\frac{y_{t+2}-y_t}{y_{t+2}+y_t}\right) + \text{abs}\left(\frac{y_{t+4}-y_{t+2}}{y_{t+4}+y_{t+2}}\right),$$

where any of these three terms equals 0 if the sign of the pre-pivot change is not different from the sign of the post-pivot change. Each of the three terms is an average of the pre- and post-pivot percent change, where the denominator in the percentage is the mean of the two years. Averaging the pre- and post-pivot percent changes is equivalent to summing the quantities—pre- and post-pivot—(difference/sum).
the trend along the way is rather different. The decline also agrees with Robert Carroll et al. (2007). The results are inconsistent with Joel Slemrod’s (1992) results for the top of the income distribution in the 1970s and early 1980s.

Figures 3.6 through 3.9 also reveal that absolute downward mobility follows a countercyclical pattern, rising during downturns and falling during recoveries (except for the recession of 1990-1991). The risk of a 25 percent drop in combined earnings is higher than that for male labor income (see Chapter Two) but lower than that for female labor income until the mid- to late 1980s. This pattern probably reflects the effect of some wives’ taking time out of the labor force for childbearing. Such interruptions increase the likelihood of a large family income drop relative to the likelihood that a man will
experience a large drop in his earnings, but the continuity of male earnings keeps the likelihood of a large income drop low relative to the likelihood a woman will experience a large drop in her earnings. At the same time, the secular increase in the number of female-headed households means that these dynamics should be less consequential over time, which is consistent with the convergence of combined head/partner downward mobility and female downward mobility levels.\footnote{When I restricted the sample to adults where the head was continuously married or single over the preceding and current calendar years, levels of downward mobility for combined head/partner earnings were very similar to levels for male labor income, but still higher than levels for female labor income.}

\begin{figure}[h]
\centering
\includegraphics[width=\textwidth]{figure37}
\caption{Percent of Adults Experiencing Two-Year Declines in Pre-Tax Household Income}
\end{figure}
Figure 3.8. Percent of Adults Experiencing Two-Year Declines in Post-Tax Household Income

Figure 3.9. Percent of Adults Experiencing Two-Year Declines in Adjusted Pre-Tax Household Income
Somewhat surprisingly, levels of downward mobility in total pre-tax family income unadjusted for family size are if anything slightly higher than in combined head/partner earnings. In results not shown, I found that pre-tax family income downward mobility was more prevalent than combined head/partner taxable income, which adds asset income to labor income. The implication is that transfer income and the income of other family members make family incomes less stable rather than providing insurance against shocks to the head’s and partner’s earnings and asset income.

Similarly, progressive taxation should also mitigate the effects of shocks to pre-tax income, since income reductions can be expected to result in a proportionally lower tax burden (Orszag, 2007). However, the pre- and post-tax risk of a 25 percent loss is very similar, implying that the tax system provides only modest insurance against large income drops. On the other hand, the risk of a $10,000 loss in income is generally lower when looking at post-tax income, but that is probably just because post-tax incomes are lower than pre-tax incomes.

The trends in absolute downward mobility are similar to the trend in downward mobility in male earnings. The likelihood of a 25 percent drop in combined labor income rose from 12 percent in the early 1970s to 17 percent in the early 2000s, while the increase for pre-tax income was from 12 to 18 percent. On the other hand, comparing cyclical troughs implies little increase since the mid-1970s, and the most recent cyclical rise in downward mobility is only somewhat higher than those in the downturns of the mid-1970s and the early 1980s. The risk of a $10,000-drop in combined labor income

57 I average the 1970 and 1972 figures and the 2002 and 2004 figures. Both of these pairs of years bracket a cyclical peak in joblessness.
rose from 15 percent to 25 percent and the risk of a $10,000 drop in pre-tax income rose from 18 to 28 percent.

Comparing my downward absolute mobility results with previous research relying on the PSID, the cyclical patterns that I find dominate any secular rise in downward mobility, contrary to the work of Hacker and Jacobs (2008), of Jacobs (2008), and of Dynan et al. (2008). My trends are basically consistent with the less-detailed results of Peter Gosselin (2008) and Gosselin and Zimmerman (2008). The general trend I find resembles what Molly Dahl et al. (2008) found in the Survey of Income and Program Participation after 1993, though I show a bigger increase and our 1980s trends are inconsistent. Finally, my results agree with those of Tom Hertz (2007), who found a rise in downward mobility over the late 1980s and early 1990s in the CPS. But while Hertz finds no change between the early 1990s and early 2000s in the CPS, I show an increase in the PSID.

For comparison, Figures 3.10 and 3.11 show trends in upward pre-tax income mobility, both unadjusted and adjusted for family size, which illustrate trends in upward mobility for the other income measures. Relative mobility changed little over the period, declining through the 1980s, and hovering between a 25 and 30 percent chance of an income increase of 25 percent or more. As in Gottschalk and Danziger (1998) and Gottschalk, McLanahan, and Sandefur (1994), upward mobility was relatively flat in the 1970s and declining in the 1980s. These trends are also consistent with Duncan, Smeeding, and Rodgers (1993). My results are consistent with Carroll et al.’s (2007) for the 1990s but not for the 1980s. They agree with my results on upward mobility in male earnings from Chapter Two.
Figure 3.10. Percent of Adults Experiencing Two-Year Increases in Pre-Tax Household Income

Figure 3.11. Percent of Adults Experiencing Two-Year Increases in Adjusted Pre-Tax Household Income
Absolute upward mobility again follows a clear cyclical pattern, with levels in the early 2000s lower than in the early 1970s. There is a downward trend from the late 1960s to the early 1980s, followed by an upward trend. These trends are basically consistent with those of Dynan et al. (2008) and Gosselin (2008), and those of Dahl et al. (2008). They also follow the trend from the early 1990s to the early 2000s of Hertz (2007), but they are inconsistent with the Hertz results from the late 1980s and early 1990s. Finally, they are consistent with the trends in upward absolute earnings mobility I found among men in Chapter Two.

Summarizing these results, then, short-term relative mobility (upward or downward) changed little over the last thirty five years. The chance of a large absolute income drop or gain rose and fell with business cycles. Large drops may have become somewhat more common, but the trend is shallow enough that it is difficult even to determine when secular declines or increases begin or end.

**Non-Directional Short-Term Mobility**

I now shift to mobility measures that do not distinguish between upward and downward mobility. These trends reflect the mix of movements in either direction that characterize volatility.

**Probability of Income Change in Either Direction.** The top two lines of Figures 3.12 through 3.15 show how the likelihood of mobility in either direction has evolved over time. The chance of moving up or down one quintile declined over the second half of the 1970s and through the 1980s but increased slightly thereafter, finishing at mid-1980s levels.
For combined labor income the likelihood of a 25 percent gain or loss rose during the 1970s, and fell by a comparable amount over the 1980s. For pre-tax family income, absolute mobility changed little over the 1970s and the first half of the 1980s but fell in the second half of the 1980s. For post-tax income, absolute mobility increased from 1981 through the mid-1980s and decreased in the late 1980s. It jumped in the early 1990s—through the mid-1990s for pre-tax and post-tax income—then changed little from the mid-1990s to 2004.

The shift in nondirectional mobility in the pre-tax income measures in the early 1990s is due to downward mobility not falling after that recession. Figures 3.6 through 3.9 indicate that the risk of a large income drop failed to rise as much during the late
1980s or to fall as much after the early-1990s recession as during past business cycles.

However, by 1998, the downward and upward mobility levels and trend mirror each other again, and the downward mobility levels appear consistent with the pre-1990 cyclical patterns.

Figure 3.13. Non-Directional Pre-Tax Income Mobility Trends
Figure 3.14. Non-Directional Post-Tax Income Mobility Trends

Figure 3.15. Non-Directional Adjusted Pre-Tax Income Mobility Trends
My estimates of trends in absolute mobility for family income and male labor income align reasonably well after 1984, but absolute mobility appears lower prior to 1984 for male earnings than it does for family income. Relative mobility also appears lower during this period for male earnings than for family income. My relative mobility findings agree with those of Gottschalk and Danziger (1998) in that we both find little change in mobility over the 1970s and declining mobility from the mid-1970s through the 1980s. The decline that I find in the 1980s and the increase that I find in the early 1990s also align with Carroll et al.’s (2007) results based on tax data. My results conflict with Maury Gittleman and Mary Joyce’s (1999) findings that relative mobility in the CPS rose during the 1970s and was higher in the early 1990s than in the late 1960s.

Dahl et al. (2008) reported that non-directional absolute mobility declined between 1985 and 1991 but was stable thereafter. I find the decline but then show an equivalent increase. In contrast, Dynan et al. (2008) report a fairly steady increase from the early 1970s to 2000.

**Intertemporal Income Association.** Figures 3.12 through 3.15 also show trends in intertemporal income association. Whether measured in terms of the Pearson correlation coefficient or the Spearman rank coefficient, intertemporal association declined over the second half of the 1970s and the 1980s, increased during the early 1990s, and was flat thereafter. The net effect was to change little over the whole period. The trend for combined head and partner earnings resembles that for male earnings in Chapter Two through the early 1990s, but then male earnings mobility increases. Through most of the early 1980s, nondirectional mobility was higher for family-level income measures than for male earnings, but over time this pattern reversed itself.
My estimates are consistent with Tom Hertz’s (2006) finding that income mobility based on the Pearson correlation coefficient rose in the CPS from 1991 to 1998 and then was unchanged in 2004.

Overall then, the four measures of non-directional mobility that I examine indicate declining mobility in the late 1970s and 1980s, increases in the early 1990s, and little change thereafter. Over the whole period, levels of mobility did not change much. There is some evidence that additional sources of income provide insurance against volatility risk. The Pearson correlation coefficient shows lower mobility in terms of pre-tax family income than in terms of combined head and partner earnings, and the risk of a 25 percent drop in post-tax income is lower than the risk of a drop in pre-tax income.

Dispersion of Income

Turning to measures based on income dispersion, either within or across people, we move further from measures of mobility and toward measures of volatility. Figures 3.16 through 3.19 display my trend estimates.

Dispersion of Income Changes. The line marked by open triangles in Figure 3.16 indicates that the standard deviation of logged head and partner labor income changes rose steadily from 1969 through the mid-1980s, but shows no consistent trend from 1984 to 2004. For pre-tax family income, volatility rose slightly in the early 1970s, then flattened out until the early 1990s when it shifted upward permanently. The post-1992 trends are basically flat. Comparing the early 1970s and early 2000s, combined head/partner earnings volatility increased by less than 25 percent, family pre-tax income
volatility rose by under one-third, and adjusted income volatility rose by about 30 percent. Volatility levels are lower than for male labor income from 1980 onwards and are significantly lower than for female labor income. Pre-tax income volatility levels are notably lower than combined labor income volatility levels, indicating that additional sources of income beyond the head’s and partner’s earnings do provide insurance against volatility. Similarly, post-tax volatility is a bit lower than pre-tax volatility.

My results are consistent with those of Dynan et al. (2008), Jacobs (2008), and Hacker and Jacobs (2008). They are also consistent with Orszag’s (2008) post-1994 findings, though not with his finding of declining volatility from 1985 to 1994, nor with Nichols and Zimmerman’s (2008) finding that volatility declined after the late 1990s.
My trend resembles that of Blundell et al. (2008) in the early 1990s but not in the 1980s. All of these studies except Orszag’s use the PSID.

**Figure 3.17. Pre-Tax Income Dispersion and Pivot Volatility Trends**

**Within-Person Income Dispersion.** The trend in the average person’s standard deviation of income over a series of years is shown in the middle line of Figures 3.16 through 3.19. The change between the comparable years of 1973 and 2000 was just under one-fifth for combined head and partner earnings and family pre-tax income. It was just 15 percent for adjusted family income. The only increase common to all four income measures occurs in the 1990s.
Figure 3.18. Post-Tax Income Dispersion and Pivot Volatility Trends

Figure 3.19. Adjusted Pre-Tax Income Dispersion and Pivot Volatility Trends
These trends are roughly similar to my corresponding male earnings volatility estimates, though income volatility levels are higher. The levels are much lower than for female labor earnings. This time it appears that the progressivity of taxes provides more insurance against volatility than either the income provided by family members or unearned income.

The increase in within-person dispersion corresponds with similar rises in practically every previous study using this concept of volatility. Gottschalk and Moffitt (2007) and Gosselin and Zimmerman (2008) each find an increase in volatility, but the timing of their increases differs somewhat from mine. Nichols and Zimmerman (2008) find that the timing of the increase depends on whether income levels or logs are examined. The less precise trends of Benjamin Keys (2008) and Craig Gundersen and James Ziliak (2008) are also consistent with my findings.

**Transitory Income Dispersion.** My last two sets of dispersion-based volatility estimates are both model-driven, and for the reasons outlined above I discount them relative to the other results in this paper. The lines marked by filled circles in Figures 3.16 through 3.19 show trends in transitory standard deviations estimated from Gottschalk and Moffitt’s variance decomposition model. According to this measure, volatility changed little between the early 1970s and the early 2000s, though the trend for combined head and partner earnings fell before rising in the early 2000s. Pre-tax income shows the biggest increase, at just 9 percent over the period.

The finding of flat income volatility runs counter to the results of Hacker and Jacobs (2008), Jacobs (2008, 2007), Gottschalk and Moffitt (2007), and Nichols and Zimmerman (2008), who all find increases. It is also contrary to my own results for male
earnings volatility. My estimates of volatility are higher for combined labor income than for male earnings, and while my volatility estimates for combined earnings start out lower than those for female earnings, the levels are similar from the mid-1980s onward. Levels of pre-tax family income volatility are, however, lower than those for either male earnings, female earnings, or combined head and partner earnings.

The trend lines marked by open circles at the bottom of the charts show estimates for transitory standard deviations from my error components model. These estimates indicate very little change from the late 1960s to the early 2000s. Only two previous studies (Blundell et al., 2008, and Blundell and Pistaferri, 2003) use error components modeling, and they feature permanent shocks in addition to transitory shocks, making them difficult to compare with my findings. The results are consistent with my male earnings volatility trends.

Summarizing the results of the dispersion-based studies, volatility appears to have risen by at most 10 percent to one third from the early 1970s to the early 2000s. In terms of levels of volatility, income from sources other than the earnings of the head and partner has a very modest insurance effect, and the progressivity of the tax system has an additional modest insurance effect.

**Pivot Volatility**

Finally, I present trends in “pivot volatility”—the mean of the percentage change in income, up and then down, for the points in a nine-year window where a reversal in trajectory may occur, averaged across individuals. Since I observe every other year in a nine-year window and need two years of data both before and after a potential pivot to
detect a change in trajectory, there are three years in which a pivot may occur. When a
reversal of trajectory does not occur at one of these pivot points, the percent change up
and then down is coded as zero. Each estimate on the lines marked by filled-in squares at
the bottom of Figures 3.16 through 3.19 shows the average adult’s pre- and post-pivot
percent change in income averaged across those three potential pivot points centered on
the estimate.58

For example, pivot volatility of combined head and partner earnings rose from
0.16 to 0.17 between 1973 and 2000, both of which were business cycle peaks. A rough
interpretation is that the change in income over the two years before or the two years after
an opportunity for a pivot was typically 16 percent of income for the window centered on
1973 compared with 17 percent for the window centered on 2000. For pre-tax family
income, the figures were 15 percent and 18 percent, adjusted or unadjusted for family
size.

This interpretation provides important context to the increases seen in the other
measures of volatility. Comparing the 1973 and 2000 values, the increase is between 10
and 20 percent, which is consistent with the increase for within-person income
dispersion. But this increase translates into minimal additional risk of volatility over
thirty years. In the early 1970s, adults could expect that a typical chance for income to
reverse trajectory would involve a rise and then a fall (or a fall and then a rise) of 15
percent over any five-year period. By the early 2000s, the typical trajectory reversal

58 The computation of this measure involves a number of averages. For a pre-pivot change of a given
individual, the change is expressed as the difference in incomes divided by the average of the two incomes
(so that increases and declines are treated symmetrically). The corresponding post-pivot change is
computed the same way. The pre- and post-pivot percent changes are averaged (with each expressed in
terms of its absolute value). This score is averaged with the scores corresponding to the individual’s other
two potential pivot points. Finally, each individual’s score is averaged with those of other individuals.
would involve 18 percent of income. The levels and trends in pivot volatility were very similar for male labor income and for family income. Levels of pivot volatility were higher for female labor income and declined over time.

**Discussion and Conclusion**

This chapter reveals a number of important facts about income instability and mobility trends. First, the various types of volatility measures yield similar results from the early 1970s to the early 2000s, ranging from a modest decline in relative mobility to little discernable change in absolute mobility and intertemporal association, to a modest increase in several dispersion-based measures of volatility and pivot volatility. A reasonable conclusion from the various trends I examine is that volatility rose by no more than a third over the thirty-year period examined. In concrete terms, this increase translates into a very small change in the magnitude of potential income trajectory reversals—the difference between income rising and falling 15 or 16 percent and income rising and falling 17 or 18 percent over five years.

These findings are similar to those for male earnings in Chapter Two, with as much or more volatility in family income as in male wages. In recent years, however, there has been more volatility in men’s total earnings, including self-employment earnings, than in family income. Family income volatility increases less than male labor income volatility does.

Measured trends are relatively insensitive to whether one looks at combined head and partner earnings, pre-tax income, or post-tax income, and adjusting incomes for family size makes little difference. Comparing levels of downward mobility and
volatility across these income definitions reveals that the income of other family
members, progressive taxation, and public safety nets provide inconsistent and modest
insurance against downward mobility and volatility.

Taken as a whole, the results of this chapter and of Chapter Two reveal little
evidence in favor of Jacob Hacker’s “Great Risk Shift” hypothesis. Families face
roughly the same risk of an income drop and the same volatility as they did in the past.
The contrast with Hacker’s conclusions can be illustrated most dramatically by
comparing his primary volatility estimates to mine. Hacker’s original claim was that
volatility had tripled over thirty years, which he has since revised downward to a 100
percent increase.59 I find that the typical family’s income reversals—up over two years
then down, or vice versa—within a nine-year window have increased only from 15 or 16
percent of income to 17 or 18 percent.

Furthermore, evidence on earnings instability from Kopczuk et al. (2009) suggests
that volatility was probably higher prior to 1960 than after 1970. Trends in various
measures of joblessness and other indicators of risk also imply that volatility was
probably no higher in 2008 than in 2002 or 2004.

To be sure, income instability and volatility levels may be too high even if they
are not increasing. But an incorrect characterization of a problem can lead to
inappropriate policy responses and ineffective political strategies by those who seek to
help workers buffeted by economic forces. An agenda organized around a “great risk
shift” may overstate the shortcomings of existing policies and may not resonate with the
middle-class voters whose political support is necessary for an expansion of public safety

I return to this theme in the concluding chapter, after exploring trends in economic instability by subgroup in Chapter Four.

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60 Winship (2007).
Chapter 4

Shaky Ground? Trends in the Risk of Large Earnings and Income Drops
Among Demographic Groups

Thus far, this dissertation has sought to describe trends in earnings and income instability, a task that has proven surprisingly difficult if the body of conflicting research produced to date is any indicator. One way the previous chapters have shed light on trends in economic instability is by analyzing trends separately for men and women. In this chapter I continue exploring subgroup differences in economic instability trends. The results provide some hints about the sources of such instability and its evolution over time.

Among the various potential sources of economic instability, and beyond the obvious first-order factors of job loss and other employment problems, sociologists might point to such factors as family change, life-cycle stage, levels of human capital, and the various structural, institutional, and interpersonal advantages or disadvantages related to ascribed race. The analyses below examine these factors, focusing on the risk of a large drop in earnings or family income.

I find that different trends across demographic groups explain little of the overall instability trends found in Chapters Two and Three. Rather, compositional shifts from relatively high instability categories to relatively low instability categories either lowered instability further than it would have declined or prevented it from rising. Specifically, the secular increase in educational attainment and the decline and delay in fertility and
relationship formation appear to be the most important factors affecting trends in economic instability.

Methods and Data

The results below build on those presented in Chapters Two and Three. Once again I use the PSID, and I keep all methodological decisions consistent with those chapters, with a few exceptions noted below. The sample is restricted to persons between the ages of 21 and 60 during the survey (which makes them 20 to 59 years old in the year for which income is reported).\(^1\) Chapter Two provides the fullest discussion of the considerations involved, while Chapter Three discusses issues specific to the income instability estimates.

Measures of Earnings, Income, and Instability

For the results in this chapter, I examine four measures of economic resources: labor income (separately for men and for women) and pretax family income (unadjusted and adjusted for family size). Labor income adds to wages and salary income a share of self-employment income and also includes earnings from a number of other sources, such as bonuses, tips, and commissions. As indicated in Chapter Two, male labor income instability appears to have increased more than male wage instability. Pretax family income aggregates the labor income of all family unit members, as well as their asset and

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\(^1\) I chose this range as a compromise between being more inclusive and keeping sample sizes large in the PSID on the one hand, and wanting to exclude students and retirees on the other. Since labor force status is reported for the current year while income is reported for the previous year, excluding students and retirees directly is problematic in the recent survey years, which were conducted only biannually. Furthermore, wives’ labor force status is unavailable in earlier years of the PSID. Note that while I exclude persons under age 20 or older than age 59, the income measures that I use that include income from all family members include income from people under age 20 or older than age 59.
transfer incomes. It does not include capital gains, lump sum payments such as lottery winnings, or in-kind transfers.

While the results in Chapter Three did not differ much regardless of whether or not I adjusted for family size, it is obvious that for some of the demographic factors I examine below—such as partnership dissolution—the amount of instability involved may depend on whether one examines family income or income per person. I adjust for family size by dividing by the square root of family size measured as of the previous survey, since respondents are asked about income received in the previous year. As discussed in Chapter Three and below, there is an inconsistency in the way that the PSID collects this information in that the survey asks about last year’s income for the current year’s family members, which is not necessarily the same as last year’s income for last year’s family members. My sensitivity tests in Chapter Three found, however, that this issue did not strongly affect the results. Since I require a measure of family size in the previous year and the PSID switched to biennial surveys after 1997, my time series for adjusted family income stops at 1996.

As in Chapter Three, for the income instability analyses, I include non-sample members who live in family units that include a sample member, in order to ensure the proper ratio between partnered adults and single adults. The earnings analyses, in contrast, are restricted to sample members of the PSID.

As discussed in the introduction, the literature on economic volatility has obscured the point that the real risk to families is that of earnings or income drops. Conditional on a given likelihood of experiencing such a drop, volatility is actually beneficial in that it amounts to recovery from drops. In this chapter, therefore, I focus on
the risk posed by large earnings or income drops, or *absolute downward short-term mobility*. I ignore downward *relative* mobility here, focusing instead on the risk of a 25 percent drop in earnings or income over a two-year period. Absolute mobility is a more direct measure of the risk of income decline than relative mobility, since the latter can remain unchanged even when income drops (if it also drops for others, as is more likely during recessions).

One departure from the previous chapters is that here I bottom-code all earnings and incomes under $1 (in 2007 dollars) to $1. This decision has the effect of including individuals who report no earnings or income for an entire year. Because I am interested in how members of different types of families fare, particularly when their family changes, and because being in a relationship can alter the decision to work or not, I want to include adults who enter or exit from work. Recoding in this way allows me to adequately compute percentage changes when a person moves from $0 to positive earnings or income while keeping the percent change when a person has no earnings in either year at zero.

As in the previous chapters, I trim the bottom and top two percent of observations (this time including the recoded $0 observations).² I adjust all incomes for inflation using the CPI-U-RS, linking it to the CPI-U for earlier years.³

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² Specifically, I trim the top and bottom 2 percent of males with positive earnings, females with positive earnings, and families with positive incomes, within age categories. The categories include male and female earners age 21-30 years, 31-40 years, 41-50 years, and 51-60 years. For the family income analyses, the age groups are based on the head of the family, and all individuals in families not trimmed are assigned the income of their family. Incomes are trimmed separately depending on the sex of the family head.

Groups Examined

Partnership and Birth Status. The PSID asks about the previous year’s income sources of the current year’s family unit members. If a couple breaks up or forms, there is a mismatch between the previous and the current years’ family unit members. The two groups will be the same only for families that experience no changes in family composition. As a result, family incomes will be mis-measured when there are family composition changes.

It is possible to get around this problem in 1968-1997, when the PSID was administered annually. To do so, I temporarily confine the sample to adults whose partnership status was the same in the survey year as in the previous year (years \( y \) and \( y-1 \)) and also consistent in years \( y-2 \) and \( y-3 \). This allows me to look at the risk of an income drop between survey years \( y-1 \) and \( y-3 \) depending on what happened to the person’s partnership status between years \( y-1 \) and \( y-2 \).

Because the PSID treats cohabiters as spouses and asks about their earnings and income if they appear in at least two rounds of the survey, I do not differentiate between spouses and multi-year partners. I used the relationship-to-head variable and the “marital status” of the head in these years, which classifies cohabiters that have lived together for two surveys as married, to group heads and wives into four initial categories: continuously partnered, continuously single, became partnered, and became single.

For the earnings analyses, I split out the “continuously married” category depending on whether or not a birth occurred between years \( y-1 \) and \( y-2 \). In particular, one might suspect that large earnings drops for women are more likely in years when they give birth. I ran into sample size problems when I tried to split other categories.
depending on whether a birth occurred. For the income analyses, I split out the “became single” category by sex, since relationship disruption could affect men and women differently. Again, sample-size considerations made it problematic to create additional categories, and trends and levels were similar for men and women for other family experiences when I experimented with additional categories. Because the income trends and levels were also similar for continuously married adults depending on whether a birth occurred, I chose not to split that group into two for the income analyses.

Because the approach I use requires annual data, the results I present by partnership status only extend through the years in which the PSID was administered annually. The requirement of having observations as far back as year $y-3$ means that my time series do not extend quite as far back as in the other analyses. The partnership status analyses cover earnings and income drops experienced in the years 1971 to 1995.4

**Age.** I consider downward mobility for four age groups: those age 20 to 29 years old in the year for which they report income, those 30 to 39 years old, those 40 to 49 years old, and those 50-59 years old. Note that students and retirees who have earnings are included in the analyses, though as noted in Chapter Two, their inclusion does not seem to strongly affect the results, and there are problems associated with trying to remove them.

**Education.** I look at separate trends for individuals with less than a high school education, high school graduates (not including GED holders but including those who go on to attend some college without getting a bachelor’s degree), and college graduates (with a bachelor’s degree or a graduate degree). From 1968 to 1990, the education

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4 I could not use the supplemental PSID Marital History File for these purposes because these files do not treat cohabiters as spouses. Cohabiters’ earnings and income were not asked about prior to the 1977 survey, but I found that no more than 1 percent of couples were cohabiters prior to 1977.
variable is based on a series of questions about the respondent’s educational history, including the highest degree they received. From 1991 to 2005, the variable is simply the highest grade completed. In the family income analyses, I use the educational attainment of the head.

**Race.** I examine trends separately for blacks and whites. Until 1985, there is no separate variable indicating the race of the wife. I assign her the race of the head in those years. Between 1973 and 1984, the head’s race was imputed based on the 1972 response, with split-off families assigned the race of the head of the family that they left. In 1985 a new race question was asked of all heads and wives, and it was asked of new heads and wives in subsequent years. Hispanicity was also determined for all heads and wives beginning in 1985. However, since the SRC sample that I use is representative only of those who were resident in the U.S. as of 1968 (and their offspring), it misses post-1968 Hispanic immigrants and their descendants. For this reason, I do not try to separate out Hispanics from other groups, and I omit the results for the “other” racial category. Also beginning in 1985, respondents could select multiple racial categories. I code sample members based on the first racial category they select. For the family income results, race is based on the race of the family head.

**Results**

Which groups face the biggest risk of a large drop in economic resources? How have relative risks changed over time? What do these patterns tell us about the source of the trends found in previous chapters? The findings here provide answers to these questions and hints for future research on trends in economic risk.
Family Composition and Change

Figures 4.1 through 4.6 present trends broken out by partnership and birth status in the risk of drops in male labor income, female labor income, pre-tax family income, and family income adjusted for family size. In these family composition charts, unlike the charts to follow, all figures are shown as three-year moving averages to smooth fluctuations arising from relatively small sample sizes. In particular, the annual sample sizes for adults who become single are quite small (under 50 cases per year, sometimes well under 50 cases). Because of the small sample sizes and the smoothing, the reader should not draw inferences about relatively small differences or changes in these charts.

Figure 4.1. Percent of Male Heads and Partners Experiencing Two-Year Declines in Labor Income of 25 Percent or More, By Partnership and Birth Status
Figure 4.2. Percent of Female Heads and Partners Experiencing Two-Year Declines in Labor Income of 25 Percent or More, By Partnership and Birth Status

Figure 4.3. Percent of Adults Experiencing Two-Year Declines in Pre-Tax Family Income of 25 Percent or More, By Partnership Status
Figure 4.4. Percent of Adults Experiencing Two-Year Declines in Pre-Tax Family Income of 25 Percent or More, By Partnership Status (Adjusted Scale)

Figure 4.5. Percent of Adults Experiencing Two-Year Declines in Family Income Adjusted for Family Size of 25 Percent or More, By Partnership Status
The most striking feature of Figure 4.1 is the high incidence of large earnings losses among men whose relationship ended between the two years earnings were recorded, particularly during the recessions of the early 1980s and 1990s. In the early-1980s recession, roughly one in three men whose relationship ended experienced an earnings drop of 25 percent or more, though sampling error may account for the extreme magnitude of the increase. In contrast, men in families with a new birth tend to have a lower incidence of earnings drops, never more than 16 percent of them from 1971 to 1995.

Both of these results are likely to strongly reflect selection. It is more likely that the employment problems that led to earnings drops also led to the breakup of relationships than that breakups led to subsequent earnings losses. Similarly, it is probably less likely that a birth reduces the risk of an earnings drop than it is that couples
tend to time their births so that they occur when the husband is economically stable. However, it is also likely to be the case that some men whose partners are expecting or who are new fathers choose to stay at their job even though they would leave if not for the new mouth to feed.

Interestingly, men who enter into a relationship during recessions have a relatively high incidence of earnings drops, even though their incidence is not particularly high during recoveries. This could reflect decisions by couples to move in together as a response to or in anticipation of the male’s employment problems, or it could indicate that when some men move to a new location to be with a new partner, they have a tougher time finding a job during a recession.\(^5\) Alternatively, it could indicate that some men take advantage of the insurance provided by a second earner to change jobs during recessions. However, if the latter explanation were salient, one would expect to see men continuously in couples also taking advantage of this opportunity, and one would expect to see men in new relationships taking advantage of the opportunity during recoveries.

The trends in Figure 4.1 reflect the cyclical pattern observed in the previous chapters, with the risk of drops higher when the economy is performing poorly. The general trend observed in Chapter Two of a rising risk of a drop through the early 1980s but no more than a small increase thereafter is reflected in the trend for continuously partnered men with no births. Continuously partnered men experiencing a birth saw their risk of an earnings drop continue to increase after the early 1980s, but the share of men who were continuously partnered and experienced a birth declined over time. Continuously single men and men who enter into relationships seem not to have faced a rising risk of earnings drops.

\(^5\) I am thankful to Christopher Jencks for suggesting this hypothesis.
Figure 4.2 presents analogous trends for women, which are dramatically different. Unsurprisingly, the highest incidence of earnings declines comes among women experiencing a birth, about 40 to 45 percent of whom see their earnings drop by at least 25 percent.

The other group who disproportionately experience large earnings drops is women who enter into a relationship, although over time the distinction between them and other women largely disappeared. This pattern likely reflects a combination of two influences. First, particularly in the earlier years, some women may have stopped working or reduced their hours upon entering into a marital or cohabiting relationship, often times in anticipation of starting a family. The sharp decline in the incidence of earnings drops among this group is consistent with this explanation, as fewer and fewer women became fulltime homemakers over this period. Second, it may be that many working women ended up having to change jobs because their entering into a relationship coincided with their partner taking advantage of a job opportunity in another geographic location.⁶

Levels of downward mobility for continuously partnered or continuously single women appear comparable to those for men whose relationship status is constant. The relatively low rates of downward mobility for women who become single compared with men who do reinforces the interpretation of the high male rates as reflecting mainly selection. Male employment problems apparently lead to relationship problems rather than vice versa. Downward mobility among women who become single is likely limited by some combination of greater labor force participation among women upon becoming

⁶ I am thankful to Christopher Jencks for also suggesting this hypothesis.
single and of pre-emptive increases in labor force participation in anticipation of becoming single.

One other notable finding in Figure 4.2 is that only one of these categories of women experiences the sizable decline in downward mobility that was observed in Chapter Two for all women. In results not shown, I found that the decline in earnings instability among women is due to a compositional shift: the number of continuously partnered women with a new birth has declined, and the number of continuously single women has increased. The shift from a high-downward-mobility category to a low-downward mobility one produces a decline in downward mobility over time.

In Figure 4.3, which shows the risk of downward mobility in family income by relationship status, I am forced to change the scale to accommodate the dramatic differences across categories. Figure 4.4 returns to the scale of Figures 4.1 and 4.2, using a second axis to show the estimates that do not fit within this range. The figures combine men and women for each category, except that they separate out men and women who become single. Clearly, the dissolution of a relationship results in a substantial shock to family income, with 60 to 80 percent of women experiencing a 25 percent drop in income and 30 to 60 percent of men. These patterns are in contrast to the stability of women’s earnings after a break up, reflecting the decline of income that occurs in moving from (often) two earners to one.

Of course, a person living alone rather than with a partner does not require two incomes, so the large income declines when relationships end may be less serious than Figure 4.3 conveys. To examine this possibility, Figure 4.5 looks at family income adjusted for family size (Figure 4.6 again uses the same scaling as Figures 4.1 and 4.2).
When incomes are adjusted for family size, men whose relationship ends experience much less instability than Figure 4.3 implies (though the risk of a 25 percent income drop has risen steadily for them over time).

In contrast, women whose relationship ends still experience a large risk of income declines, even adjusting for family size. These patterns presumably reflect at least two factors. First, it remains the case that male earnings typically make up a larger share of family income than female earnings do, so when relationships end, the economic loss will be felt more by women. Second, when a couple has children, the female partner is more likely than the male partner to have custody of the children. Because any child support contributed by a mother’s former partner will be notably less than his income was, the woman’s family income is likely to fall more than her family size. From the early 1970s to the early 2000s, 45 to 80 percent of women whose relationship ends experienced a 25 percent drop in size-adjusted family income. The risk of a large income drop has declined sharply over time, though it remains far above the risk for other groups.

Figures 4.3 and 4.5 show a corresponding but less dramatic difference in the risk of an income drop for adults who enter into relationships. If family income is not adjusted for family size, the risk is less than 10 percent for these adults. It is rare for the addition of a potential earner to lead to a large drop in income. However, when incomes are adjusted for family size, the risk increases to roughly 10 to 20 percent, similar to that faced by continuously single or partnered adults. This risk is still low, however. In 1995, the last year for which I can break these categories out, just one in six working-age adults who did not experience the breakup of a relationship or a birth saw their family-size-adjusted incomes drop by 25 percent or more, and this risk had not increased appreciably
over time. Indeed the risk only increased appreciably among men whose relationship ended and among partners experiencing a birth.

One final pattern worth noting is the difference between Figures 4.3 and 4.5 in instability patterns for partners experiencing a birth. Before adjusting income for family size, the risk of a large income drop for continuously partnered adults appears unaffected by births. This surprising result may be due to several factors. Many couples experiencing a birth may already have a non-working wife, in which case there would be no change in income. For other couples, the time taken off by the wife after delivery may be short enough that the resulting income loss is less than 25 percent. And there is probably some selection in that many couples are likely to time their births so that they are doing well financially in the year the child is born or in subsequent years. Note, however, that after adjusting for family size, Figure 4.5 shows that the addition of a new family member clearly translates into a higher risk that economic resources per person will drop by 25 percent or more.

Age

In addition to family changes, economic instability is also likely to change over the course of adults’ working lives, as they gain experience, settle into a career path, and approach retirement. Figures 4.7 through 4.10 show trends in the risk of a large income loss for four age groups – persons in their 20s, 30s, 40s, and 50s. Looking at male earnings instability first, adults age 20 to 49 have similar instability levels and trends, as shown in Figure 4.7. Older men, however, appear to have a higher risk of downward earnings mobility after the early-1980s recession. This pattern could simply reflect a
decline in age at retirement, or a more gradual transition into retirement, in which older men reduce their work hours prior to dropping out of the labor force entirely. Or it could indicate a secular shift upward in the instability faced by older workers. The risk of an earnings drop does not appear to have increased after the early 1980s for any age group.

Figure 4.7. Percent of Male Heads and Partners Experiencing Two-Year Declines in Labor Income of 25 Percent or More, By Age

Figure 4.8 indicates that trends for women are different, with little cyclicality, young women experiencing a declining risk of an income drop, and older women experiencing a rising risk. The pattern among younger women reflects declining fertility as well as increasingly delayed marriage and fertility.

Turning to family income, Figure 4.9 shows that income instability is higher for younger and older families than for families with a head age 30 to 49. Families tend to have the same trends in instability regardless of the head’s age. Instability has risen a bit among families with a head under age 40 since the early 1980s, although without
knowing what happened after 2004, it is difficult to say this conclusively. Finally, Figure 4.10 indicates that after adjusting for family size, young adults experience the highest levels of family income instability, with levels 5 percentage points higher than adults age 30-49. Older adults’ instability levels and trends look much like those of adults 30-49 until the mid-1980s, after which levels increased to those of young adults.

![Figure 4.8. Percent of Female Heads and Partners Experiencing Two-Year Declines in Labor Income of 25 Percent or More, By Age](image)

**Education**

One of the primary developments in the American economy in recent decades is the rise of the wage premium for college graduates, as the supply of college graduates failed to keep pace with demand. Figures 4.11 through 4.13 examine whether the growing disparity in earnings between those with relatively little education and college graduates translated into differences in economic instability.
Figure 4.9. Percent of Adults Experiencing Two-Year Declines in Pre-Tax Family Income of 25 Percent or More, By Age

Figure 4.10. Percent of Adults Experiencing Two-Year Declines in Family Income Adjusted for Family Size of 25 Percent or More, By Age
Figure 4.11 shows that even after the early 1980s, the risk of a large earnings drop rose for high school dropouts, high school graduates, and college graduates alike. Among college graduates, it rose from around 10 percent in the early 1970s, to around 12 percent in the early 1980s, to 15 or 16 percent in the early 2000s (comparing similar points in the business cycle). Instability was higher among men with less education, particularly during the recessions of the mid-1970s and the early 1980s (when it briefly exceeded 30 percent). In fact, the reason that the risk of a large income drop stayed fairly constant among all men after the early 1980s is that there was a large compositional shift in the population from men with less than a high school degree to men who have a college degree. This shift drew men into more stable educational categories.

Figure 4.11. Percent of Male Heads and Partners Experiencing Two-Year Declines in Labor Income of 25 Percent or More, By Educational Attainment
Figure 4.12. Percent of Female Heads and Partners Experiencing Two-Year Declines in Labor Income of 25 Percent or More, By Educational Attainment

Figure 4.13. Percent of Adults Experiencing Two-Year Declines in Pre-Tax Family Income of 25 Percent or More, By Educational Attainment
Women also benefited from this shift. Figure 4.12 shows that the risk of a large earnings drop probably rose after the early 1980s among the least-educated women and fell only among college graduates. But the number of women with college degrees increased over the period, putting more and more of them into a category with lower earnings instability and a flat or declining trend. Particularly in recent years, the least educated women have experienced greater risks of downward mobility than other women.

Finally, family income instability trends by the education of the head are shown in Figure 4.13 (I omit results for family-size-adjusted income, which gives similar results). Clearly instability is higher the lower education is, and this disparity is stronger during recessions. The risk of a large income drop increased after the early 1980s only among families headed by a college graduate, rising from about 13 to 14 percent of adults to about 15 to 19 percent. But because these families had relatively low instability compared with other families, the increase in college graduation among family heads meant that overall income instability did not increase.

Race

The last demographic breakdown I examine is race. The PSID began in the wake of the civil rights gains of the mid-1960s and the racial tensions that followed. Did the experience of African Americans improve relative to whites? Figure 4.14 shows the familiar pattern of increases in male earnings instability through the early 1980s, with little secular trend after that. The estimates for blacks are imprecise, but instability is
generally higher among black than white men. There is no evidence that blacks gained (or lost) ground relative to whites.

Figure 4.14. Percent of Male Heads and Partners Experiencing Two-Year Declines in Labor Income of 25 Percent or More, By Race

Turning to female earnings, Figure 4.15 implies that the decline in instability among women through the mid-1980s was mainly among non-whites, while the trend among white women was relatively flat. Again, there is little change after the mid-1980s for either group. The estimates for black women are too imprecise to determine whether levels differ for blacks and whites. Apparently the racial differences are not as great for women as they are for men, at least from the mid-1970s onward. The early decline in instability among black women may have been a result of civil rights legislation, but that interpretation requires an explanation for why black men did not experience a similar improvement.
Figure 4.15. Percent of Female Heads and Partners Experiencing Two-Year Declines in Labor Income of 25 Percent or More, By Race

Figure 4.16. Percent of Adults Experiencing Two-Year Declines in Pre-Tax Family Income of 25 Percent or More, By Race
Finally, Figure 4.16 indicates little secular change in family income instability for either whites or blacks (family-size-adjusted results are similar). Families with a black head experience higher instability rates than those headed by a white adult. This finding is perhaps surprising in light of the very different rates of marriage among African Americans compared with whites.

Discussion and Conclusion

The trends observed in Chapters Two and Three obscure a number of interesting differences across demographic groups. First, levels of economic instability differ across groups. Earnings instability is especially high among men and women who do not have a high school diploma, among black men, among women under 30, among women who experience a birth, among men and women who enter a relationship (though less so in recent years among women, and only during recessions among men), and especially among men whose relationship ends. Over time, several of the patterns for women have weakened, however, as they have increasingly come to delay marriage and childbearing and made up some of the ground separating them from men in their wage levels.

Earnings instability is lower than average among male college graduates, white men, men over 50 years old, and men with a new birth. It seems likely that these patterns largely reflect the importance of the skills that workers bring to the labor force, or selection that is related to these skills. In addition, it is possible that greater instability among black men reflects discrimination, inadequate access to job networks, or other factors beyond human capital levels. Greater instability among older men almost certainly reflects a good deal of planned reduction in labor supply.
In terms of family income, groups with relatively high instability include partners experiencing a birth, women exiting a relationship, younger families, and families with heads who are black or high school dropouts. Clearly family change and economic opportunities are both important factors determining female instability levels. Presumably, much family change is anticipated, in which case the resulting instability may be mitigated by precautionary savings or other steps.

As for trends, Chapters Two and Three found that in the 1970s earnings instability rose for men and fell for women, while family income instability showed little secular trend. In contrast, I found increasing male earnings instability after the early 1980s for men with low, medium, and high education levels. Thus the flatness of male earnings instability after the early 1980s is a consequence of rising educational attainment on the part of men, which masks rising instability conditional on education. Earnings instability also rose after the early 1980s among men experiencing a break up and among partnered men experiencing the birth of a child. On the other hand, it declined among men older than 50 years old.

Among women, the risk of a large earnings drop increased after the early-1980s only among those without a high school diploma. It declined among women under 30 and among college graduates. The decline in earnings volatility among women in Chapter Two is largely a function of compositional shifts. Over time, fewer women were partners who gave birth, and more were single. Furthermore, the educational levels of women also rose over time. Both of these compositional changes left more women in relatively low-instability demographic groups.
Finally, turning to family income instability, families headed by a college graduate saw increases in instability over time, as did families with a head under age 40. Both men and women in relationships who experience a birth also suffered rising income instability (adjusting for family size), and men experiencing a breakup did as well. On the other hand, income instability declined among women who experienced a breakup.

In one sense, the demographic variables considered in this chapter explain very little of the trends observed in the preceding chapters. In results not shown, I compared the risk of a drop in male earnings, female earnings, and family-size-adjusted family income for all adults and for those age 30 and up with at least a high school education, no births, and who were continuously partnered or single. While levels of instability were consistently lower for the latter group than for all adults, the trends were scarcely affected.

The effect of demographic characteristics on instability trends lies in the importance of compositional shifts – increases in educational attainment and, among women, declines in fertility and relationship formation. These shifts moved adults into more stable demographics, which caused economic instability to be lower than it otherwise would have been.

Having shown that economic instability has not increased over time, I conclude with a review of trends in other indicators of economic risk that are largely consistent with the instability trends. The conclusion also addresses the related question of whether economic insecurity has grown before delving into the politics and policy of economic insecurity and risk.
Chapter 5

Conclusion: Economic Risk, Policy, and Politics

The foregoing chapters have shown that economic risk, at least as evidenced by trends in earnings and income instability, has risen far less than Hacker and others have claimed. Using this yardstick, the economy has become riskier for typical American families, but not a lot riskier. Below I will show that this is also true of other indicators of economic risk.

Economic risk need not be rising to justify public policies to improve the economic security of Americans—it just needs to be “too high”. However, those concerned about economic security must squarely face two sets of facts that challenge beliefs that are widespread on the political left. First, levels of risk are not rising as much and are not as high as many progressives believe. Second, economic anxiety is not as deeply felt as they believe. Perhaps partly as a consequence, there is only modest appetite among Americans for grand new social insurance programs.

I present evidence for both of these contentions below before concluding with a programmatic approach to economic risk that addresses these political constraints. My proposal essentially consists of voluntary federally-sponsored offerings in the vein of conventional employee benefits. I illustrate this approach by outlining new health care reform and retirement savings proposals.
The Great Risk Shift Reconsidered

Hacker and others who believe in a “risk shift” offer two rationales for their agenda: a policy one and a political one. The policy rationale asserts that the magnitude of the problems facing the middle class is large enough that its members are only marginally more secure than the poor. Therefore, only universal policies based on social insurance are up to the task of alleviating the economic insecurity of the middle class. Few observers have elaborated this belief as articulately as Hacker does in the concluding paragraph of *The Great Risk Shift*:

> The philosophy behind the reforms that we need is one of constructive change guided by an abiding spirit, the spirit of shared fate. Today, when our fates are often joined more in fear than hope, when our society often seems riven by political and social divisions, it’s hard to remember how much we all have in common when it comes to our economic hopes and values. Indeed, we are more linked than ever, because the Great Risk Shift has increasingly reached into the lives of all Americans….What the ever-present risk of falling from grace reminds us of is that, in a very real sense, all of us are in this together. The Great Risk Shift is not “their” problem; it is our problem, and it is ours to fix.¹

> The problem with this rationale for expansive (and expensive) policy is that its proponents overstate both the depth of the economic problems facing the middle class and the extent to which these problems have worsened. As the preceding chapters have shown, the extent to which economic instability has increased has been greatly overstated. The same is true of other indicators of supposed risk. A careful reading of Hacker’s work finds little in the way of trend data related to the problems he discusses. I begin with an indicator that on first glance appears to bolster his case.

**Risk of Joblessness.** The question of whether the job stability has fallen has been the subject of heated debate for the past twenty years, and the evidence remains

¹ Hacker (2008).
somewhat ambiguous. Much of the early research suffered from data inconsistencies. The most careful research tended to find flat or modestly declining job stability—measured as current tenure or job retention rates—from the mid-1970s to the early 1990s.² But David Neumark and his colleagues (1999) found a modest decline in job stability over the first half of the 1990s,³ David A. Jaeger and Ann Huff Stevens (1999) found that the share of men age 40 and older who had under ten years of tenure began to rise in the mid-1980s, and two studies using the National Longitudinal Survey of Youth found bigger declines in job stability among men under 40.⁴ Leora Friedberg and Michael Owyang (2004) also found that average job tenure among men declined from 1983 to 1998 for groups defined based on how long they have been out of school. Among women, tenure declined among younger workers, but may have increased among older workers.⁵

Two contributions to a new volume on the risk of job loss convey the current state of debate on the question. Stevens (2009) finds that the average man near age 60 in 1969 had held his longest job for 22 years.⁶ That was the same duration for the average man near age 60 in 2004. In addition, over half had worked on a job for 20 years or more in both years. However, tenure did decline from 1975 or 1980 to 2004. An important limitation of this study is that it reflects the experiences of men at the end of their careers. Since there is evidence of a secular decline in job stability since 1985 or 1990, today’s

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² Diebold, Neumark, and Polsky (1997); Jaeger and Stevens (1999); Gottschalk and Moffitt (1999); Farber (1998); Bansak and Raphael (1998); Fitzgerald (1999).
³ Neumark, Polsky, and Hansen (1999).
⁴ Jaeger and Stevens (1999); Monks and Pizer (1997); Bernhardt et al. (1999).
⁵ Friedberg and Owyang (2004).
younger workers will have had different experiences by the time they are 60 than Stevens’s sample. However, if most workers hold their longest job near the ends of their careers, then one would expect that recent changes would have shown up in Stevens’s results.

Henry Farber’s (2009) paper in the same volume—the latest in a series of papers he has written over the course of a decade—finds that mean tenure by age declined among men between the 1973-79 period and the 2000-06 period.7 The share of men with at least ten years or at least twenty years of tenure also declined. Among women, tenure increased. However, Farber’s results may not adequately account for rising educational attainment—which reduces career length on the front end—or declining age at retirement (or semi-retirement), which reduced career length on the back end.8

In a more reliable set of results, Farber finds that the share of men in the private sector with less than one year of tenure rose modestly from 1973 to 2006 (from about 9.5 percent to about 12.5 percent), while the share of women with less than one year of tenure fell. His evidence implies that the increase among men was confined to workers age 30-64 and was particularly large among men in their thirties.

It appears that job turnover has become more common in the past two decades. But if our concern is economic risk, this trend is not very informative. Job stability can reflect the preferences of either employers or employees. Low job stability can reflect

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7 Farber (2008).

8 In some analyses, Farber estimates mean tenure or the probability of long tenure for each year using age and education main effects in a regression model, but because the relevance of education to tenure varies with age, the model should interact age and education. Farber (2007) estimates mean tenure for each birth cohort by regressing it on age for four separate educational categories, but it is not possible to translate the cohort effects into year effects. The cohort estimates reflect the calendar years that a birth cohort is in the data. Stevens (2009) finds that cohort trends based on Farber’s method follow trends based on hers rather closely, and she estimates that the predicted average year of longest tenure at age 60 will fall to 20 years by 2011.
poor economic conditions (in which case lower job security leads to shorter tenure) or strong economic conditions, where strong labor demand leads workers to switch to better jobs. Data from the Bureau of Labor Statistics Job Openings and Labor Turnover Survey indicate that depending on the point in the business cycle, 40 to 60 percent of job separations not due to retirement, death, or transfers involve voluntary separations chosen by the employee.\(^9\) Rather than trends in job stability, what is really of interest are trends in job security. The fact that earnings instability rose only modestly among men and fell among women suggests that rising job instability has probably involved a lot of voluntary rather than involuntary job separation. Trends in a wide variety of other statistics also support this conclusion.

Start with unemployment. The official unemployment rate has shown a secular decline since the recession of the early 1980s, as shown by the “U3” line in Figure 5.1. There are a number of issues that make the unemployment rate an imperfect measure of joblessness, but the Bureau of Labor Statistics has tracked additional measures intended to address these shortcomings. As Figure 5.1 shows, the trend in joblessness is unaffected by the definition used.\(^{10}\) Note, in particular, that the trend in the share of the labor force that is unemployed because they were laid off (U2)—rather than leaving a job or entering the labor force after being out of it—follows the same trend.

Other measures of joblessness also fail to substantiate the great risk shift hypothesis. Job losers as a percent of the unemployed have shown no secular trend since the early 1980s (Figure 5.2). “Expected job loss”, which adjusts cross-sectional figures

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to account for the fact that some have just started a jobless spell and for the greater likelihood that the long-term jobless will be sampled at a point in time, has been stable or declined since the early 1980s, both for all of the unemployed and for job losers.\footnote{Valletta (2005).}

**Figure 5.1. Trends in Joblessness**

![Figure 5.1. Trends in Joblessness](image)


Steven Davis (2008) summarizes research that consistently finds that flow rates from employment to unemployment have declined since the late 1970s or early 1980s.\footnote{Davis (2008).} In addition, job destruction rates have declined since the 1980s or 1990s, and one time
series assembled by Davis and his colleagues shows that they have declined in manufacturing since the early 1960s.\textsuperscript{13}

![Figure 5.2. Job Losers as a Percent of the Unemployed](image-url)

Jay Stewart (2002) of the Bureau of Labor Statistics found little change in job separation—departure from a job—between the mid-1970s and the early 2000s.\textsuperscript{14} However, that trend masked an increase in job separations that produced spells of unemployment under two weeks and a decline in separations that produced longer spells of unemployment. Stewart interprets these findings as evidence that more job separations involve direct employer-to-employer transitions than in the past—that is, that they are largely employee-driven. Robert E. Hall (2005) found that the job-finding rate (new

\textsuperscript{13} Davis et al. (2006).

\textsuperscript{14} Stewart (2002).
hires divided by the sum of the unemployed, those marginally attached to the labor force, and discouraged workers) has seen a secular increase since the early 1980s.\textsuperscript{15}

On the other hand, median weeks unemployed has risen, comparing trough to trough, since 1989, though it rose even more in the late 1960s and 1970s (Figure 5.3). The share of the unemployed who have been out of work for 27 weeks or more has also risen fairly steadily since the late 1960s (Figure 5.4).

\begin{figure}
\centering
\includegraphics[width=\textwidth]{Figure5.3.png}
\caption{Median Weeks Unemployed}
\end{figure}


But again, context is everything. Unemployment itself fell after the early 1980s, and the two measures in Figures 5.3 and 5.4 apply only to the small share of the labor

\textsuperscript{15} Hall (2005).
force that is unemployed. So in 2008, when the median duration of unemployment among the unemployed was 9.4 weeks, median duration of unemployment among all adults in the labor force was 0 weeks (as was the 25th and the 10th percentile of unemployment duration!). The share of the unemployed that had been out of work for 27 weeks or more was 20 percent, but the share of the labor force unemployed for 27 weeks or more was 1 percent (Figure 5.4). Neither of these figures—applied to the entire labor force—grew over time.

**Figure 5.4. Percent Unemployed for 27 Weeks or More**


Finally, the share of the employed who are part-time workers but would like to be full-time and the share of part-time workers who would like to be full-time also has declined (Figure 5.5).
Note that a number of these indicators show that joblessness has increased, trough to trough, since 2000, but that hardly constitutes solid evidence of a permanent risk shift. At least through 2008, levels of joblessness were well within the range seen in the past forty years, and there is no reason to think that they will remain permanently elevated after the current recession ends. What stands out is how consistent the evidence is that the economy has been working as well as it ever has for most of the American workforce over the years. Again, that does not mean it is working well enough or that things have not worsened for some categories of workers, but there has been no major shift in employment risk in recent decades.
Health Care Risk. The share of the population without health coverage has risen over the past twenty years (see Figure 5.6), from 12 percent in 1988 to 15 percent in 2007, according to the most-cited government source (the Current Population Survey). The decline in private coverage among the non-elderly has been steeper, dropping from 79 percent in 1980 to 67 percent in 2007 (Figure 5.7). The shallower trend for the decline in any insurance coverage reflects the increasing role played by public coverage, which in turn reflects some combination of government responding to greater risk in the private sector and crowd-out of private coverage by public coverage.

![Figure 5.6. Percent of Population with no Health Insurance, 1984-2007](http://www.census.gov/hhes/www/hlthins/historic/hlthin05/hihistt2.html; http://www.census.gov/hhes/www/hlthins/historic/hihistt2.xls) and the National Health Interview Survey, 1984 and 1997-2007 (http://www.cdc.gov/nchs/data/sr_10/sr10_162.pdf; http://www.cdc.gov/nchs/data/nhis/earlyrelease/insur200806.htm). CPS figures are the percent uninsured for the entire previous year, though most experts believe some respondents report their current insurance status. There is a break in the series in 1999 – I adjust the pre-2000 figures so that the percent difference remains the same going back in time. NHIS figures are the percent currently uninsured.
However, these changes must be put in historical perspective. First, coverage levels remain high by historical standards. It is possible to build a consistent health coverage time series extending back to 1940. Doing so produces the trend shown in Figure 5.8, which makes clear that what has happened is that progress in increasing the number of Americans with insurance coverage has hit a floor. The changes of the past twenty to twenty-five years are minuscule compared with past levels of non-coverage. It would certainly be better if Figure 5.8 continued to show a downward path, but the risk of
lacking health coverage has clearly not increased dramatically, and the current period remains substantially less risky than the entire 1940-1980 period. It is true that medical expenses have become far more costly over time relative to typical incomes, but catastrophic costs to the uninsured have always been catastrophic.

**Figure 5.8. Percent of Population Currently Without Health Insurance, 1940-2007**

A second point of context is provided by national health expenditure data, which fleshes out what “health insurance coverage” includes and how that has changed over time. Figure 5.9 reveals that private insurance—which, unlike out-of-pocket expenditures, pools risk—has come to cover a greater share of private expenditures on all forms of personal health care, with dramatic increases for most categories of care. In most cases, coverage leveled off only after the early 1990s. The notable exception is prescription drug coverage. And there were expansions of coverage even within these categories. Coverage of outpatient mental health services, for instance, became much more generous over this period.\(^\text{16}\)

\textbf{Figure 5.9. Percent of Private Expenditures Paid by Private Insurance, 1960-2007}

\vspace{1cm}


\(^{16}\) Blumenthal et al. (1988), p. 58-60.
So even while the share of Americans without private health coverage was declining, coverage grew much more generous for those who retained it. And the public sector caught most of those who otherwise would have become uninsured, particularly through expansions in Medicaid in the late 1980s and the creation of the SCHIP program in 1997. But isn’t it still the case that employers have shifted an increasing share of costs of private coverage onto their employees? Given the growing generosity of benefits, it would be unsurprising to find such a trend. But apparently that is not what has happened.

Figures 5.10 and 5.11 show the distribution of health expenditures on health services and supplies by source of spending for 1987 and 2007 respectively (the earliest and latest years for which I could find data). Private employer contributions to insurance constituted 18 percent of health expenditures in 1987 and 19 percent in 2007. Adding in employer contributions to Medicare, the total was 23 percent in both years. Expenditures by households, on the other hand, fell from 41 percent of health expenditures to 31 percent. Most economists would argue that employer expenditures for their employees’ health insurance come out of employee compensation, in which case it is appropriate to allocate employer expenditures to households. In that case, the share of health spending by households fell from 64 percent to 54 percent (or from 69 percent to 61 percent if public employer expenditures are included).
Figure 5.10. National Expenditures on Health Services and Supplies, 1987

- Private Employer Contributions to Private Insurance Premiums and Medicare: 23%
- Households (in yellow): 41%
- Medicare: 17%
- Public Programs: 39%
- Government: 43%

Figure 5.11. National Expenditures on Health Services and Supplies, 2007

- Private Employer Contributions to Private Insurance Premiums and Medicare: 23%
- Households (in yellow): 31%
- Medicare: 19%
- Public Programs: 44%
- Government: 50%

Perhaps most persuasively, household out of pocket spending fell from 23 percent of national health spending in 1987 to 13 percent in 2007. Public and private employees paid 23 percent of employer-based insurance premiums in 1987 and 27 percent in 2007 (not shown in Figures 5.10 and 5.11). If premiums for individually purchased insurance are included, the change was from 29 percent to 31 percent. That is the extent of the risk shift.

The picture that emerges from these figures is quite different from that painted by Hacker. Over the past 50 years, consumers have demanded more and more from their health insurance. Employers have responded in two ways: by expanding the scope of coverage and by picking up more of their employees’ health care expenses, essentially taking the extra cost out of their wages, and by reducing the share of the workforce that is covered by health insurance. The expansion of public insurance has mitigated the problem of health coverage in one sense, by filling in for private insurance, but it is possible that the expansion has also contributed to private coverage declines (the phenomenon of “crowd out”). When it comes to health care, the issue is clearly the dramatic increase in the cost of health care. The extent to which this increase simply reflects rising consumer demand, valuable innovation, and increased well-being rather than being a national crisis is beyond the scope of my research into whether there has been a Great Risk Shift.

**Retirement Security Risk.** Participation of private labor force workers in employer-sponsored retirement plans was about 45 percent in both 1975 and 1995, and it has since crept upward.\(^\text{17}\) The share covered by a traditional defined benefit pension

\(^{17}\) U.S. Department of Labor (1999), Table E4, available at http://www.dol.gov/ebsha/publications/bullet1995/e_4.htm; U.S. Department of Labor (2008), Table E4,
shrank by half over this period, but coverage in defined contribution plans more than compensated. Furthermore, Figure 5.12 shows that real private employer costs for their workers’ retirement and savings rose by about 20 percent between 1991 and 2008, a rate of increase that mirrored the growth in employers’ spending on health insurance. If there has been a risk shift in retirement savings, it cannot be about coverage declines or employer stinginess.

Figure 5.12. Private Employers' Hourly Costs for Workers' Retirement and Health Insurance


The conventional argument for a risk shift in retirement security is that the growth of defined contribution plans, such as 401(k) plans, and withering of defined benefit plans has shifted risks from employers to employees. This is almost surely true, but two caveats are rarely acknowledged. First, with additional risk comes additional opportunity, and a rational assessment of increases in risk must take into account both costs and benefits. Second, there are also risks associated with defined benefit plans. Hacker presents a remarkably one-sided view of the risks inherent in defined contribution plans relative to traditional pensions. To illustrate the trade-offs involved, suppose you had to choose between these two retirement savings schemes at the start of your career. Which would seem riskier to you?

(1) You will work for one or more employers throughout your career, each of whom will lower your take-home pay by some amount in order to have enough money to pay you upon retirement. At age 65, you can begin receiving monthly retirement benefits for the rest of your life. You will not be able to touch these benefits until then. The amount of these benefits will be calculated as follows: for each job that you work at least six years, you will receive one percent of the earnings you made in the last year you were there, times the number of years you were there. If you work a job for, say, only three years, you will receive just 40 percent of that amount, and you won’t receive anything unless you stayed there for at least two years. Therefore, you will have incentives to remain in your jobs for a number of years, and you will be at risk of losing some or all of your
benefits if you are laid off as a relatively new hire. Under this system, it will be particularly important that you achieve your peak earnings at the end of your career in a job where you are entitled to benefits, since inflation will erode the value of your benefits if your maximum earnings come relatively early in your career. There is no guarantee that you will have more upon retirement than you would have had if your pay had not been lowered to pay for your retirement and you had been able to invest the additional amount on your own. Finally, if one of your employers declares bankruptcy, you may not receive the full amount that you thought you were entitled to. You will, of course, have Social Security benefits as a safety net.

(2) You will work for one or more employers throughout your career. Your paychecks will be bigger than under Option (1) because your employers will have a less open-ended commitment to your retirement savings. You may use that additional income to save for retirement, but if you prefer spending it (or need to spend it) on other things, you are free to do so. If you choose to save for retirement, you will contribute part of each paycheck to an individual retirement account, chosen from among several offered by your employer. Your employer will match any contribution you choose to make, up to three percent of your salary. As early as age 59 but no later than age 70, you will receive the assets that have accumulated over your career, including the original contributions and the returns they
generated. Only those employer contributions from jobs where you stayed four years will be included – those from other jobs where you left before that will go back to your employer, along with the returns they generated. Otherwise, your retirement assets remain yours and continue accumulating returns when you switch jobs or if you go into business for yourself. Furthermore, you will have the ability to tap into your retirement savings for limited purposes, such as emergencies, though sometimes subject to heavy tax penalties. You will be responsible for saving enough for your retirement and for managing your retirement assets, which may erode in value depending on the assets in which you invest and broader economic conditions. In general, you will face a trade-off between financial risk and the size of potential returns. Finally, when you eventually receive your assets upon retiring, you will have no guarantee that they will support you throughout the rest of your life, and it will be up to you to manage them to ensure you maintain your living standards as best you can. You will, of course, have Social Security benefits as a safety net.

These two options correspond, respectively to idealized defined benefit and defined contribution systems. It should be clear that there are risks to both systems, and equally clear that it is difficult or impossible to choose between the two options without knowing something about the probabilities of doing well or poorly under each.\(^\text{18}\) That is precisely what James Poterba and his colleagues attempt to ascertain in a recent paper.\(^\text{19}\)

\(^\text{18}\) For a discussion of the risks associated with each type of retirement plan, see Bodie et. al (1988).

\(^\text{19}\) Poterba et al. (2007).
By simulating the distribution of wealth accruing to retiring workers after career-long participation in defined benefit and defined contribution plans, using contribution rates, plan parameters, financial returns, and pension formulas derived from data and typical practices, the authors estimate how many people are likely to fare better or worse under different plan types. They find that a worker with median earnings for his or her education will be better off under a defined contribution system than under a system based on private sector defined benefit plans. Indeed, the highest-earning 90 percent of all workers will tend to fare better under a defined contribution plan if it includes investment in stocks and stocks perform at their historical average. Of course, not everyone at every point in time will do better under a defined contribution plan than a defined benefit plan—witness the evaporation of wealth that occurred in the portfolios of near-retirees who were heavily invested in stocks (or real estate) in 2008 and 2009. But over the long run, Poterba’s results imply that most people will be better off under a defined contribution regime than a defined benefit regime.

Nevertheless, some who believe that the economy has become riskier claim that the shift from defined benefit plans to defined contribution plans dooms large swaths of the population to inadequate retirement savings.\textsuperscript{20} The research on this question comes to inconsistent conclusions, in particular the subset of studies defining saving adequacy in terms of how post-retirement living standards compare to pre-retirement ones.\textsuperscript{21} The

\textsuperscript{20} For example Christian Weller, then with the Economic Policy Institute, claimed in 2002 that, “The average American household has virtually no chance to reach an adequate retirement savings in the next 50 years.” (Dugan, 2002).

\textsuperscript{21} Bernheim (1993, 1997); Bernheim et al. (2000); Gale (1997); Moore and Mitchell (2000); Gustman and Steinmeier (1998); Yuh et al. (1998); Montalto (2001); Wolff (2002); Warshawsky and Ameriks (2000); Center for Retirement Research (2006); Court, Farrell, and Forsyth (2007); Scholz, Seshadri, and Khittrakun (2006); Eggen, Gale, and Uccello (1999, 2005). See Congressional Budget Office (2003) for a review.
disagreement is largely due to problems associated with the difficulty of forecasting income, savings, and consumption paths, including measuring pre-retirement consumption adequately, determining what level of retirement income will allow retirees to live at the level their pre-retirement income afforded, and modeling rates of return, life expectancies, and changes in savings behavior as retirement approaches and in the event of shocks to income and wealth. John Karl Scholz and his colleagues have argued that the very standard of maintaining pre-retirement consumption is an inappropriate one for economic models, because large families will consume more during their working lives (by spending on their children) but will require less when those children have left home.\textsuperscript{22} While there are reasons to question this assertion, its plausibility highlights the difficulties inherent in defining “adequacy” of savings. In general, the research producing the most pessimistic findings has been persuasively challenged on methodological grounds\textsuperscript{23}, while the research that most adequately incorporates economic uncertainty into its modeling has produced relatively optimistic findings.

More to the point, almost none of these studies examine whether adequacy of retirement savings has declined.\textsuperscript{24} On the other hand, studies that compare current retirees or near-retirees with past ones, or that compare savings of current workers to those in the past, tend to find that recent cohorts are as prepared for retirement and as

\textsuperscript{22} Scholz, Seshadri, and Khitatrakun (2006).


\textsuperscript{24} Center for Retirement Research (2006) is an exception, finding that the share of the population at risk of inadequate savings rose between the 1980s and 2004. See Scholz and Seshadri (2008), however, for a critique of their methods.
wealthy or wealthier in retirement. The Congressional Budget Office (2004), for instance, showed that wealth-to-income ratios were similar from 1983 to 2001 for 35- to 44-year-olds and 45- to 54-year-olds. Robert Haveman and his colleagues (2007) found that annuitized net wealth was 60 percent higher for married couples retiring in the 1990s than it was for couples retiring in the early 1980s, and one-third higher for single retirees.

To be sure, there are problems associated with our movement toward a defined contribution system, including relatively low participation rates, high cash-out rates when workers change jobs, and insufficient annuitization options upon retirement. But these problems are fundamentally different from those created by a risk shift in which employers are increasingly dumping workers into hopelessly inadequate private accounts that are doomed to leave retirees destitute.

**Risk of Indebtedness.** Hacker declares that “[p]ersonal bankruptcy has gone from a rare occurrence to a relatively common one,” citing figures from Elizabeth Warren showing an increase from 290,000 in 1980 to over 2 million in 2005. Figure 5.13 confirms a sharp rise in bankruptcy rates, expressed as a share of households. However, a peak rate of 1.45 percent in 2003 can hardly be called “a relatively common” occurrence.

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25 Congressional Budget Office (2004); Haveman et al. (2007); Manchester, Weaver, and Whitman (2007); Butrica, Iams, and Smith (2007).

26 I could not find figures for cumulative risks of bankruptcy over longer periods. In results not shown, I found that the risk of experiencing a 25 percent income drop between two years at least once rose from 16 percent from 1995-96 (only one opportunity for a drop) to 37 percent from 1986-96 (ten opportunities for a drop). If these figures are any guide, then even at a constant annual bankruptcy rate of 1.45 percent the rate over a decade would be well under one in twenty.
A problem with the bankruptcy figures is that they can change for reasons other than hardship. For example, bankruptcy filings rose in the lead-up to the 2005 federal legislation that tightened eligibility for bankruptcy, in anticipation of the change, and it plummeted after the bill was implemented. Furthermore, the stigma associated with bankruptcy can change over time, as can the intensity of marketing by firms specializing in bankruptcy litigation. It is more appropriate, for looking at trends in over-indebtedness, to consider measures that are less affected by such issues.
One such measure is the typical amount of debt payments when expressed as a percentage of income, which is estimated in the Federal Reserve Board’s Survey of Consumer Finances. The median ratio of debt payments to family income (among persons with debt) increased from 15.3 percent of income to 18.0 from 1989 to 2004—a very modest rise. This conclusion also appears to hold if one looks at the least wealthy quarter of the population. The median among the least-wealthy quartile increased from 10.1 to 13.0. These modest increases occurred during a period in which personal bankruptcies rose dramatically, suggesting that the increase in bankruptcy rates was at least partly about an increase in the willingness to file for bankruptcy conditional on debt.27

Additionally, the share of debtors with a ratio of debt payments to income that exceeded 40 percent grew from 10.0 to 12.2 percent between 1989 and 2004 (8.1 to 10.6 among the least-wealthy quartile). The share of debtors with payments that were delinquent by 60 days or more increased between 1989 and 2004, from 7.3 to 8.9 percent and from 17.6 to 22.9 percent among the least-wealthy quartile.

Home foreclosure rates? Hacker noted the rate had increased fivefold since the early 1970s. Figure 5.14 shows that things have gotten much worse with the bursting of the housing bubble. But even at the end of 2008, when the rate was at an all-time high, just 3.3 percent of mortgages were in the foreclosure process. That translates into about 2.3 percent of homes, since three in ten homeowners has no mortgage.28 Again, the point is not that we should be unconcerned that one in fifty homeowners are losing their homes.


28 Bucks et al. (2006).
The point is that an increase in the risk of losing one’s home from one in 300 to one in fifty does not constitute a fundamental risk shift that will provoke anxiety in the typical American family.29

Figure 5.14. Mortgage Foreclosure Rates, 1970-2008

Economic Insecurity and Political Preferences

It should be clear that the “risk shift” as a rationale for new universal, expansive social insurance policies is one that rests on a foundation of sand. The second rationale

29 If the annual foreclosure rate remained as high as its current levels over the long-run, that would imply a true risk shift, since an annual one in thirty chance over, say, thirty years would produce a 100 percent chance of eventual foreclosure. If the evidence for income drops noted in footnote 26 is any guide, and if foreclosure rates settle at 1 percent annually, the implied risk of foreclosure over thirty years is probably well under one in ten.
for these policies is a political one. Advocates argue that their proposals will be advantageous to the party that adopts them, because they accord with the preferences of voters for a more ambitious federal role in solving economic problems. Hacker again is eloquent on this point, asserting that “the rising risk to the economic well-being of families posed by the post-1970s transformation of our economy, including the substantial decline of employer and governmental policies of insurance” is “at the heart of the increasingly negative verdict of most Americans -- and, yes, married Americans with degrees and kids -- about the direction of the economy”. It is also, according to Hacker,

the issue with the greatest potential to unify Americans across lines of class, race and education behind a new economic agenda that provides a basic foundation of security so as to guarantee true opportunity. It is also, not incidentally, the issue with the greatest potential to create a powerful new coalition behind the Democratic Party”.

Hacker argues that economic insecurity is rising by pointing to a handful of survey results, the most prominent of which are from International Survey Research (ISR, now part of Towers Perrin). Its figures—from American workers at large employers—showed that the share of workers “frequently concerned” about “being laid off” rose from 14 percent at the height of the early 1980s recessions, to 46 percent in 1996, when the economy was poised to take off, before falling to a still-high 35 percent in 2005.

Figure 5.15 shows the best figures for a nationally representative group of workers that I could find. It shows two series going back to the mid-1970s, both of which consistently asked employed adults how likely it was that they would lose their jobs in the next year. One series, the General Social Survey, comes from the National Opinion Research Center and includes substantial involvement from academic

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30 Hacker (2006b).
researchers. The other comes from the Gallup Organization. In contrast to the 14 percent to 35 percent growth in job insecurity that the ISR data indicate, the GSS shows the fraction who think it “very” or “fairly” likely that they will “lose their job or be laid off” falling from a high of 15 percent in the early 1980s to 10 percent in the mid-2000s, and Gallup shows a decline from 18 percent to 10 percent.

Figure 5.15 indexes the GSS and Gallup figures to their values as of the winter of 1976-1977 and does the same for the official monthly unemployment rate. Clearly these two measures have the expected relationship to the unemployment rate, particularly before the recession of the early 1990s. After this recession, however, a gap opens up between the series and unemployment trends, indicating that job insecurity did not fall as much in the past twenty years as employment trends would have predicted.

What accounts for this shift in insecurity? The recession of 1990-1991 was marked by widespread commentary about downsizing and the unprecedented effect it was having on middle-class workers rather than being confined to less-skilled labor. In retrospect, the media attention devoted to the “white-collar recession” was far out of proportion to the downturn’s impact on middle-class workers. But there is suggestive evidence that it had a permanent impact on the psychological state of American workers—a great risk-perception shift.

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31 The recession had mild effects on unemployment compared with past recessions, both for blue-collar and white-collar workers, and the effect was stronger on blue-collar workers than white-collar workers. But relative to past recessions, the 1990-91 recession was somewhat less mild for white-collar workers. See Gardner (1994).
Figure 5.15. Adults Saying It Is Very or Fairly Likely They Will Lose Their Job in the Next Year

Source: General Social Survey – authors calculations; Gallup – Roper Center, iPoll (http://www.ropercenter.uconn.edu/data_access/ipoll/ipoll.html); Unemployment Rate – Bureau of Labor Statistics (http://www.bls.gov/cps/), not seasonally adjusted. NBER recessions are shown as shaded bars.
Indeed, to the extent that the ISR figures can be reconciled with the GSS and Gallup figures, it appears that the effect of media hand-wringing about the middle class may play a determinative role.\textsuperscript{32} Figure 5.16 displays the ISR trend, and the unemployment rate, indexed to their 1979 values (the first year of the ISR data). The two series do not track each other very well, in contrast to the evidence in Figure 5.15. However, after the early 1990s recession, an enormous gap opens up between the two series.

Figure 5.16 also shows a third time series, plotted on a second axis that is scaled so that the levels of the three series are comparable—the number of articles from the New York Times and the Washington Post in each year that mention either “downsizing” or “offshoring”.\textsuperscript{33} This time series follows the ISR trend remarkably well, spiking upward during and immediately after the early 1990s recession and failing to decline to its previous levels as the economy strengthened.

Hacker and others are correct that insecurity has risen, even though the shift appears not to reflect actual changes in economic risk, and even if the trends in nationally representative measures indicate no rise in insecurity (see Figure 5.15).\textsuperscript{34} In another sense, however, they misdiagnose the public mood. Regardless of trends, economic anxieties are not as deep as claimed by Hacker and others on the political left.

\textsuperscript{32} I was unable to obtain detailed methodological information about the ISR surveys. They appear to be confined to workers at firms with at least 500 employees (Kessler, 2000). Such firms employ roughly half of workers in the U.S. (see http://www.census.gov/epcd/www/smallbus.html).

\textsuperscript{33} These figures were obtained from a LexisNexis Academic search.

\textsuperscript{34} In addition, Jacobs and Newman (2008) show that trends in consumer sentiment and in financial decline over the previous year show declining insecurity over time. The latter indicator tracks monthly unemployment remarkably well.
Figure 5.16. Percent of Workers Frequently Concerned About Being Laid Off, and Newspaper Coverage of Downsizing and Off-Shoring

Surveys that ask Americans about their economic anxieties are ubiquitous. It is difficult to determine what to make of responses to such questions, however, without context. One way to provide context is to look at trends, as in Figures 5.15 and 5.16. Another is to compare the anxieties expressed about economic concerns with other anxieties. For example, Harris Interactive conducted two surveys for the American Psychological Association in 2008 on the stress experienced by adults in the U.S. While financial and work-related issues are the biggest sources of stress, “job stability” ranked lower than all other sources except personal safety, including lower than family responsibilities and relationships.35

Since the early 1990s, the Kaiser Family Foundation has been asking survey respondents how concerned they are over different potential problems. While they have documented significant economic insecurity, anxieties unrelated to economics are at least as widespread. On the one hand, as the financial crisis was escalating in October of 2008, nearly half of adults said they were “very worried” about their income “keeping up with rising prices”. Nearly four in ten said they were very worried about the possibility of “having to pay more for health care or health insurance”. Roughly one in three were very worried about “losing their savings in the stock market” and also about “not being able to afford needed health care”.36


On the other hand, two-thirds of parents were also very concerned that their children are “exposed to too much inappropriate content” in entertainment and mass media.\textsuperscript{37} Two-thirds of adults in a 1996 Kaiser survey said they were very concerned about their “taxes getting too high”.\textsuperscript{38} Nearly half were very concerned about “government interfering too much in [their] life”. About three in ten adults are very concerned about illegal immigration.\textsuperscript{39} One in three is very concerned about the possibility of a dangerous error while they fly on U.S. airlines.\textsuperscript{40} Two in ten are very worried about being the victim of a violent crime, and the same fraction is very worried about being the victim of a terrorist attack.\textsuperscript{41} Those figures are broadly similar to the 25 to 30 percent who said at the end of 2008 that they were very worried about losing their job or their health insurance.\textsuperscript{42}

Furthermore, there is little evidence from these surveys that concern over economic issues has increased. In mid-1996, nearly six in ten adults said they were “very concerned” about their “income keeping up with the cost of living”, and fear about inflation seems to have declined between 1996 and 2004.\textsuperscript{43} The share of adults worried about having to pay more for health care or health insurance increased from early 2000 to

\begin{itemize}
\item \textsuperscript{37} See \url{http://www.kff.org/entmedia/upload/7638.pdf}.
\item \textsuperscript{38} See \url{http://www.kff.org/kaiserpolls/loader.cfm?url=/commonspot/security/getfile.cfm&PageID=14464}.
\item \textsuperscript{39} See \url{http://www.kff.org/kaiserpolls/upload/Immigration-in-America-Toplines.pdf}.
\item \textsuperscript{40} See \url{http://www.kff.org/spotlight/mederrors/upload/Spotlight_Jan06_MedErrors.pdf}.
\item \textsuperscript{41} See \url{http://www.kff.org/kaiserpolls/upload/7831.pdf}.
\item \textsuperscript{42} Ibid.
\item \textsuperscript{43} See \url{http://www.kff.org/kaiserpolls/loader.cfm?url=/commonspot/security/getfile.cfm&PageID=14464} and \url{http://www.kff.org/kaiserpolls/loader.cfm?url=/commonspot/security/getfile.cfm&PageID=32224}.
\end{itemize}
Concern about being able to afford needed health care apparently declined from 1996 to early 2004 and declined more from 2004 to late 2008. As recently as April of 2008, just one in six adults were very worried about losing their savings in the stock market, so the high current levels are a reaction to the financial crisis.45

The “I’m-OK-They’re-Not” phenomenon also provides crucial context for interpreting the many polling results that indicate concern about the economy.46 There is a long-standing tendency for survey respondents to say that while they themselves—or their family, their children, their child’s school, or their political representatives—are succeeding, the rest of the country is doing badly.

Three recent prominent polls about the state of the economy illustrate this phenomenon. In a New York Times/CBS News poll from early April of 2009, with unemployment nearing thirty-year highs, half of respondents said the condition of the national economy was “very bad”, and only 11 percent said it was very good or fairly good. In contrast, two-thirds of these same respondents said the financial situation of their household was very good or fairly good, and only 12 percent said very bad.47

In the Allstate/National Journal Heartland Monitor Poll that same month, 30 percent said that most Americans were doing poorly at “managing the economic and financial opportunities and risks they face”, and another 54 percent rated most Americans

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44 Ibid.


“fair”. However, 56 percent of respondents said that they themselves were doing a good job or excellent job, with only 5 percent saying a poor job.48

Finally, a poll by Greenberg Quinlan Rosner Research and Public Opinion Strategies for The Pew Charitable Trusts Economic Mobility Project in January and February 2009 found that 55 percent of respondents thought that “people in this country today” are not very much in control or not at all in control of their personal economic situation. But three in four felt they themselves were very much or somewhat in control.49

David Kusnet, Lawrence Mishel, and Ruy Teixeira (2006) have argued that people accurately describe the (poor) economic situation of others while misstating their own situation.50 They offer two hypotheses for why this might be the case. First, they argue that many survey respondents cannot bring themselves to admit that they are not doing well, and so they lie (to the pollster, or even to themselves). Second, they argue that people just cannot see that they are doing relatively poorly because that would require them to compare themselves to people of the same age in the previous generation. Instead, they compare themselves to how they were doing a few years ago, and because people generally earn higher wages as they age and gain skills and experience, they find that they are doing relatively well.

Kusnet and his colleagues offer no evidence for their hypotheses. In contrast David Whitman (1998) offers a range of examples not confined to economics that show

49 See http://www.economicmobility.org/assets/pdfs/Poll_Questionnaire.pdf. As a disclosure, I am employed by the Economic Mobility Project and was centrally involved in the planning and analysis of the poll. None of the views expressed in this thesis should be taken as representing those of the Project or of The Pew Charitable Trusts.
50 Kusnet et al. (2006).
how people systematically evaluate conditions in the larger society pessimistically compared with how they evaluate their own life. In addition, for the second hypothesis to be true, one has to believe that when people judge how they themselves are doing, they look back a few years rather than a generation, but when they judge how everyone else is doing, they look back a generation rather than a few years. That is possible, but not obvious. The principle of Occam’s Razor would argue that these findings simply reflect the fact that people have more and better information about how they themselves are doing than about how others are doing. When media accounts emphasize the economic risks facing people the result is surface-level insecurity based on a belief that one has somehow escaped the calamity that is befalling the rest of the country.

Off Center Reconsidered

Because many progressive advocates of a great risk shift are misinformed about economic trends and the nature of economic anxiety, a market arose during the Bush years for books that explained how the political right had managed the success that it enjoyed despite economic conditions that should have favored the left.51 One of the most influential of these was Jacob Hacker and Paul Pierson’s book Off Center: The Republican Revolution and the Erosion of American Democracy, which appeared in 2004 before The Great Risk Shift. Hacker and Pierson argued that what pundits and political observers called political polarization had been a one-sided affair.52 Republican activists and legislators had grown more conservative, but Democratic activists and legislators had

51 Frank (2004); Lakoff (2004); Armstrong and Moulitsas (2006); Westen (2007); Krugman (2007); Bartels (2008).

52 Hacker and Pierson (2005).
not grown more liberal (and had even moved to the right themselves in some regards). Along with this shift, Republicans had developed effective strategies to move public policy further rightward than the typical voter preferred prior to the 2006 election.

Since the rightward shift of Republicans occurred during a period in which Hacker and Pierson showed the distribution of self-identified ideology had not changed, the implication was that the electorate was being deprived of the more-progressive policies that it desired. While the Democratic Congress that emerged from the 2006 elections found it politically difficult to enact its agenda during the last two years of the Bush presidency, the Democratic presidential victory in 2008 and expanded majorities in Congress have renewed optimism that the preferences of that typical voter will finally be satisfied.

Unfortunately for these optimists, Hacker and Pierson’s analysis is riddled with problems.

**Do Americans Strongly Support Additional Government Spending?** Hacker and Pierson’s argument that the contemporary Republican agenda is out of line with voter preferences relies most heavily on the case of the Bush tax cuts. While a small majority of Americans continued to favor tax cuts in 2004, Hacker and Pierson emphasized that this support cannot be meaningfully interpreted without posing tax cuts against other priorities such as spending increases.

Hacker and Pierson argue persuasively that the polling data show that tax cuts were a relatively low priority for Americans during the Bush years and that Americans generally preferred that any tax cuts go to less wealthy citizens. But they do not make a compelling case that public preferences in favor of alternative priorities were strong.
Thus, it is not clear that Bush’s success in selling the tax cuts depended on the success of the Administration efforts to obscure the tax cuts’ size and distribution.

Consider first the question of the public’s understanding of the Bush tax cuts. According to 2004 American National Election Study data, six in ten Americans favored the Bush tax cuts as of 2004. Forty-two percent strongly approved while 26 percent strongly disapproved. Even after these tax cuts, 40 percent of Americans thought their own taxes were too high in 2004, compared with just 5 percent who thought they were too low. Half thought the taxes of the poor were too high, compared with 7 percent who thought they were too low. On the other hand, 63 percent thought that the taxes of the rich were too low, versus just 11 percent who thought they were too high. Without even considering trade-offs, then, there was little strong opposition to the Bush tax cuts, and there was little support for a tax increase unless it fell on the rich.

Perhaps supporters of the tax cuts were misinformed, victims of the disinformation campaign that Hacker and Pierson document? This appears not to be the case. Among those Americans who favored the Bush tax cuts, only one-third believed the average worker’s taxes had gone down. Among supporters of the tax cuts, half thought the taxes of the rich were too low. Apparently, many Americans must have supported the tax cuts simply because they generally supported Bush’s agenda, for the effect they thought the cuts had on the economy, or for other reasons.53

As Hacker and Pierson note, polling responses related to budget priorities are much more meaningful when respondents are forced to choose between lower taxes and greater spending (or lower budget deficits). ANES respondents are explicitly asked to

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53 This conclusion is contrary to that of Bartels (2008, Chapter 6), who argues that political ignorance mostly explains support for the tax cuts.
choose between the budget deficit, domestic spending, and taxes. This series was asked in both 1996 and 2004. In 1996, 48 percent of Americans supported raising taxes, increasing the deficit, or both in order to increase domestic spending. The corresponding figures for tax cuts and deficit reduction were 46 percent and 44 percent, respectively. Spending increases, tax cuts, and deficit reduction had basically the same levels of support when Americans were confronted with the tradeoffs involved.

More importantly, what I will call “strong support” for spending increases was even smaller. Just 16 percent of voters indicated a clear preference for spending increases when pitted against tax cuts or reducing the budget deficit. That is, they not only favored spending increases over both the alternatives, they also opposed spending cuts to cut taxes or reduce the deficit. This compares with 4 percent who preferred tax cuts and 6 percent who preferred reducing the deficit. Strong support for spending increases exceeded both, but it was itself still very low.

In 2004, the budgetary outlook and tax code were dramatically different than in 1996, courtesy of the Bush tax cuts. With taxes lower and the deficit much bigger, we would expect support for tax cuts to have shrunk and support for deficit reduction to have grown. Instead, 53 percent supported cutting taxes, an increase of 7 points, and just 38 percent supported deficit reduction (a decline of 6 points). Meanwhile, two-thirds of Americans in 2004 supported spending increases over at least one of these two alternatives—a 20 point increase.

Strong support for both spending increases and tax cuts increased as well, but it remained confined to small minorities. Just 23 percent of Americans in 2004 were strong supporters of spending increases, compared with 9 percent who strongly preferred tax
cuts and 3 percent who strongly preferred deficit reduction. Once again, spending garnered the most support, but only a quarter of Americans strongly favored it.

The weakness of Americans’ political preferences has important implications for Hacker and Pierson’s contention that policies under Bush were “off center”. Take support for greater public education spending, which stood at 74 percent among Americans in 2004. If one counts as supporters only those who also strongly favored domestic spending increases over tax cuts and deficit reduction, then only 21 percent of Americans strongly favored increases in public education spending. One can be a strong supporter of education spending without being a strong supporter of domestic spending in general, but presumably ANES respondents have their own pet projects in mind when thinking about “domestic spending”. The implication is that even a policy area as popular as public education – which had the most support out of eleven policy areas mentioned in the survey – has surprisingly soft support.

Similarly, half of Americans preferred that government provide more services (with greater spending) rather than fewer services (with lower spending), compared with just 24 percent who preferred the opposite. But if one counts only those who are strong supporters of increasing spending, the percent supporting more services and spending drops to 15 percent. Preferences for greater spending are apparently broad but shallow.

Hacker and Pierson argue that the Bush tax cuts should not have passed if policy is based on the views of the median voter. As noted, just 4 percent of Americans consistently favored tax cuts over the alternatives in 1996. But if one looks at the share that would have supported tax cuts over either deficit reduction or spending increases (but not necessarily over both), that figure increased to 46 percent, which wasn’t
statistically different from support for spending increases over one of the two alternatives. Unfortunately the tradeoff questions were not asked in the 2000 ANES.

**Has the Republican Congress Swung Far to the Right of the Typical Voter?**

Not only do Americans have only weakly held preferences for additional spending or tax cuts, but as Hacker and Pierson note and as I will show below, their ideological dispositions appear not to have changed much over time. If the Republican Congressional Caucus grew more conservative over time, that would support Hacker and Pierson’s contention that by 2004 the caucus was “off center”—far to the right of the median voter.

To argue their case, Hacker and Pierson turned to scores created by Keith Poole and Howard Rosenthal that put Members of Congress past and present on a common scale measuring ideological position.\(^{54}\) Hacker and Pierson report that the polarization of Congress between the early 1970s and the early 2000s was almost entirely due to growing extremism among Republicans. Democratic legislators had not moved nearly as far from the center. Because of the increasing conservatism of Republicans, Congress was, in the early 2000s, far to the right of the median voter, who had not grown more conservative over time. But Hacker and Pierson’s account is flawed.

Consider the Senate.\(^{55}\) Poole and Rosenthal’s scores, using every vote by every Member of every Congress through the 108th Congress (which ran from 2003 to 2004), indicate that the “center” as of 2003-04 was typified by northeastern Republicans such as

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\(^{54}\) See [http://www.voteview.com/](http://www.voteview.com/).

\(^{55}\) Following their recent book, *Polarized America* (McCarty, Poole, and Rosenthal, 2006), I use scores on Poole and Rosenthal’s first DW-NOMINATE dimension (for details, see [http://polarizedamerica.com/](http://polarizedamerica.com/) and [http://www.voteview.com](http://www.voteview.com)). Hacker and Pierson report using “d nominate” scores, but these are only constructed through the 99th Congress, so I am inclined to believe that they too used the first DW-NOMINATE dimension scores.
Lincoln Chafee, now-Independent Jim Jeffords, and William Cohen; Arlen Specter (now a Democrat); and by red-state Democrats such as Ben Nelson and John Breaux. In 1971-72, the median Senator had a score of -0.056, equivalent to Ben Nelson’s score in 2003-04. By 2003-04, the median Senator had a score of 0.061, equivalent to Arlen Specter in 2003-04.

This small change in the median of the Senate as a whole only hints at the fact that, as Hacker and Pierson claim, Republican Senators did move farther ideologically than Democratic Senators. The evidence that Hacker and Pierson presented describes how the median in one year compared with then-recent Senators’ scores. In the early 1970s, according to Hacker and Pierson, the median Republican Senator lay “significantly to the left of current GOP maverick John McCain of Arizona—around where conservative Democrat Zell Miller of Georgia stood” [where the references to McCain and Miller are to their 2003-04 scores, italics in the original]. The median Republican Senator’s score then “doubled” by the early 2000s so that it sat “just shy of the ultraconservative position of Senator Rick Santorum.”

This claim raises a technical issue. The Poole-Rosenthal scores are not ratio scales with a meaningful zero point. The distance between 0.2 and 0.4 is supposed to be the same as that between 1.2 and 1.4, but 1.2 is not “six times as conservative” as 0.2, because a score of 0 does not indicate the complete absence of conservatism. The zero point is completely arbitrary. The doubling from 0.2 to 0.4 would become an increase of just 50 percent if we added 0.2 to all of the scores (from 0.4 to 0.6). We cannot know

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56 These descriptions do not quite reflect what the Poole-Rosenthal scores show. The median Republican Senator’s score in 1971-72 was equidistant between McCain in 2003-04 and Miller in 2003-04, not closer to Miller, and it was just as closer to McCain than the median Republican Senator’s score in 2003-04 was to Santorum.
whether Republican Senators grew twice as conservative between the early 1970s and the early 2000s. Indeed, the phrase “twice as conservative” has no obvious meaning.

Hacker and Pierson’s interpretation of these results is an even bigger problem. Rather than the Republican Party drifting ever rightward (the whole time increasingly “off center”), if the Democratic Party was “off center” in the early 1970s, then the movement among Republicans could be interpreted as a restoration of an equilibrium reflecting voter preferences. This is exactly what appears to have happened.

First of all, the medians for the 2003-04 Senate were 0.379 and -0.381 for Republicans and Democrats – essentially identical. That means that after this great rightward shift by Republicans, the parties were equally “extreme” by historical standards. Furthermore, the median Democratic Senator in 1971-72 wasn’t much less extreme than the median Senator from either party in 2003-04.

Second, at least in terms of self-identification, the ideological distribution of Americans was unchanged over this period, with roughly twice as many people calling themselves conservative as calling themselves liberal.57

Taking these facts together – a rightward shift by Republican legislators, an end state where Democrats and Republicans are equally “extreme”, and an ideological distribution among voters that was static over the period (and right-leaning) – the conclusion that best fits is that the Democratic Congress of 1971-72 was off center rather than the Republican Congress of 2003-04. The median Republican became more extreme over time, but that was because Congress became more representative of the

57 Hacker and Pierson (2004), page 38. Hacker and Pierson cite ANES data. According to Gallup data showing self-identified ideology, the breakdown among Americans as a whole in 2004 was roughly 20 percent liberal, 40 percent moderate, and 40 percent conservative (Wave 2 of the June Poll, Question D10). In 1972, it was 25 percent, 34 percent, and 37 percent (Poll 851, Question 14).
The story on the House side is much the same, except that the median Republican was a bit more “extreme” than the median Democrat by 2003-04 (although no more extreme than the median Democrat was in 1971-72).

The Bush Administration and the Republican Congress may have used various tactics in order to pass an agenda that lacked strong support. But they were not “off center” if that phrase is taken to mean that their agenda was outside the bounds of what the public supported.\footnote{Hacker’s and Pierson’s argument that Republican activists grew more extreme while Democratic activists became less so is also problematic. To support their claims, Hacker and Pierson began by defining an activist as someone who self-identifies as a Democrat or a Republican and who participated in three out of five election-related activities asked about in the American National Election Studies. They measured ideology using a combination of two “thermometer” items – one of which asks respondents how warm or cold they feel toward liberals and one inquiring about conservatives. The resulting measure ranged from 0 (extremely warm toward liberals and extremely cold toward conservatives) to 97 (extremely cold toward liberals and extremely warm toward conservatives). To determine how far activists drift from the center, they compared the activist scores on this index to the scores for independent voters. Hacker and Pierson plotted the average distance from independents for Republican and Democratic activists and then “smoothed” the trends by imposing curves to describe them. The graph showed that Republican activists were more extreme than Democratic activists to begin with, that they became more conservative over time, and that after becoming more liberal, Democratic activists tacked back toward the center. But several of these conclusion disappear when the data points are examined rather than the curve being fit to them, when 2004 data are added, when activists are compared not against Independents (which can produce an increase in extremism without activists’ ideology changing) but against the numerical center of the scale, or when self-identified ideology is used rather than a measure that conflates ideology with tolerance toward other ideologies. They also treat their measure as if it had a true zero point as discussed above, indicating inappropriately that Republican extremism doubled over time. When these issues are addressed, Republican and Democratic activists are equally extreme in several years (including 2004) and there is a clear increase in extremism among Democratic activists.} Or more specifically, where Republicans succeeded, their agenda was not out of bounds. Hacker and Pierson downplayed the extent to which Republicans had to reach out to the center in what they did or did not favor. Education spending, for instance, increased more under Bush than under Clinton, in a nod to “compassionate conservatism”.\footnote{See \url{http://www.gpoaccess.gov/usbudget/fy09/pdf/hist.pdf}.} Furthermore, where Republicans truly moved off center, they failed, as with Social Security privatization.
Finally, voters cannot have representation that aligns perfectly with their preferences on all issues; they must optimize as best they can. There is little discussion of national security policies and “values issues” in *Off Center*, but the Republican advantage in these areas is stronger (or was before the Iraq war in the case of national security) than in tax and spending policy.\(^6\) It may be that much of the public has stronger preferences in areas of Republican strength than in tax and spending policy, and that they are willing to tolerate tax and spending priorities that they support only weakly if at all.

The conclusion that voters have only weak preferences when it comes to taxes, spending, and deficits contains good and bad news for those who advocate a broader risk-mitigation agenda. On the one hand, it means that ambitious spending programs are palatable to a majority of Americans when taxes are as low as they were in 2004. Progressives can also be heartened that tax cuts commanded strong support only from a tiny fraction of Americans in 1996 and 2004. On the other hand, domestic spending does not have very strong support – and certainly not as strong as many conventional polling questions imply. And ambitious domestic programs were not supported over tax cuts or deficit reduction by a majority of Americans in 1996, when taxes were higher and the budget deficit was lower. Furthermore, tax cuts were also palatable to a bare majority in 2004, even after the Bush tax cuts. We do not have similar questions for 2000, when the budget was briefly in surplus and taxes were somewhat lower than in 1996, but a majority of Americans indicated in 2004 that they approved of the Bush tax cuts.

\(^6\) Galston and Kamarck (2005).
Citizen Benefits: Formulating Policy to Address Economic Risk in the Context of Political Constraints

As of the spring of 2009, when the fate of the American and world economies still remained unclear, those promoting the risk-shift hypothesis were having an I-told-you-so moment. But however long or deep the recession turns out to be, the economic damage that occurs to American families will bear only tangentially on the risk-shift hypothesis. The current recession, caused by the rapid proliferation of financial securities subject to minimal regulation, is likely to be *sui generis*, and the policies most urgently needed to prevent another recurrence involve not changes to the social contract, but regulation of financial institutions and investments. Prior to the recession, many economists talked of a “Great Moderation” in describing the modern macroeconomy, with its relatively tame recessions and modest recoveries. Clearly there were unappreciated developments working at cross-purposes to the Great Moderation, but there is no reason to think that we have entered a new period of Great Volatility.

Recession aside, the evidence on trends in economic risk indicates that for the typical American family, economic life is only marginally riskier today than in the past. Trends in economic anxiety show that anxiety has increased since the early-1990s recession. The persistence of the “I’m OK, They’re Not” phenomenon and expressions of non-economic anxiety in polling implies that people nevertheless perceive less risk to themselves than they believe others face. And the evidence on spending, taxation, and deficit reduction preferences suggests that permanently raising federal spending to levels above where they stood at the end of the Clinton years will be politically difficult (particularly with budget deficits as high as they currently are).
All of this evidence, taken together, suggests that successfully implementing universal, high-price-tag insurance programs to protect against economic risk will be extraordinarily difficult. Hacker has proposed influential health insurance and “universal insurance” policies. His health insurance proposal has been estimated to cost the federal government $49.3 billion annually. However, a careful reading of the tables in the cost estimate memo indicates that it would also add $47.2 billion in costs to employers who would otherwise not offer coverage but would be required to pay taxes to the federal government under the proposal, bringing the cost to almost $100 billion.

Other estimates floating around Washington health policy circles in spring of 2009 implied a cost more on the order of $130 billion or even $170 billion. Hacker estimates that his universal insurance proposal would cost another $35 billion annually.

Policies aimed at reducing risk or the consequences of risk must be politically feasible. They also ought to have a healthy respect for the risk of unforeseen consequences. As an alternative to Hacker’s approach, federal policy could move in the direction of a policy regime I call “citizen benefits”.

As Hacker has documented, the United States is uniquely dependent on employers to provide non-cash benefits of various types, particularly health insurance and retirement savings. The system made great sense in earlier decades, when lifetime employment

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64 Hacker (2006c).
65 Hacker (2002).
was the expectation by employer and employee alike and when few married women worked.

Today, however, employer benefits have a number of disadvantages as well. Employer-provided benefits burden employers with risks that other rich countries have socialized, potentially hurting America’s international competitiveness. They impede job mobility, leave workers vulnerable when they are out of work, and lead workers to make poor savings decisions upon job loss. They also discourage cost containment in areas such as health care by obscuring trade-offs faced by employees and insulating consumers from their choices.

Under a program of citizen benefits, Americans would have access to a range of benefits sponsored not by their employer, but by the federal government. Just as workers today have employer-sponsored retirement plans and health insurance policies, all citizens would have access to government-sponsored benefits as well. Contrary to many proposals offered by the political left, citizen benefits would be entirely voluntary – there would be no mandates imposed on individuals or employers—and the current employer-based benefit system would continue to exist alongside it. There would be subsidies to the disadvantaged, but they would be much more transparent than under today’s arrangements or those in other policy proposals. And tradeoffs between greater pay, richer benefits, different benefit mixes, and higher costs would be much more transparent than today. A federal safety net would continue to exist to catch those who fall through the cracks of the citizen benefit system.

To see how such a system would work, consider health care insurance. My proposal essentially consists of reorganizing the health insurance system so that subsidies
are directed to citizens rather than employers and so that existing employer-sponsored
insurance is supplemented by a federally-sponsored insurance pool that anyone could join
(but that no one would have to join). Most individuals would end up with health
coverage as an employer benefit or, analogously, as a citizen benefit.

To switch to this system of citizen benefits, two large tax expenditures focused on
employer-sponsored insurance—the deductibility of insurance-related health expenses by
employers and the exclusion of health benefits from individual income taxation—would
be repealed. The primary effect of this change would be for many employers providing
health insurance benefits to drop coverage and instead increase the pay of their workers.
The additional earnings would be subject to individual income and payroll taxes, but
changes in tax rates and brackets could be enacted simultaneously if the additional
revenue generated were large enough to pay for the additional costs of the program. In
other words, employees currently receiving employer-sponsored health insurance would
end up subsidizing the currently uninsured, but they would still see after-tax wage
increases, leaving them with higher take-home pay with which they could purchase
health insurance.

Or not purchase it. Individuals would not be required to buy health insurance, and
employers would not be required to provide it or be taxed to support citizen benefits
(though the repeal of the employer deduction would raise corporate income taxes).
Employers could choose to sponsor coverage, largely subject to existing federal and state
law, in which case premiums would be deducted from the paychecks of participating
employees. Insurers providing employer-sponsored health insurance would also operate
as they do today, with no additional regulation of rate-setting, benefit requirements, or
underwriting. For workers currently covered by an employer, then, very little would change under the proposal if their employer continued offering coverage, except that cross-subsidies and trade-offs would become more transparent.

For those whose employer dropped coverage, who never had employer coverage, or who simply prefer an alternative to their employer’s coverage, a federal insurance pool would be created that anyone could join. The pool would resemble the Federal Employees Health Benefits Program (FEHBP) in that private insurers would offer different coverage options and benefit packages and compete on price. To create a large initial population, the FEHBP would be folded into the new pool. The new pool’s plans would feature guaranteed issue and renewal, and unlike most FEHBP plans, its plans would be community rated. These features of the pool would mean that the only way to ensure that premiums remained sufficiently low for the healthy to participate would be to subsidize participating insurers for the costly enrollees they cover. Therefore, a federal reinsurance program would pay 75 percent of all individual claims over $30,000. Insurers would still compete on price in covering these high-cost individuals, both because they would be responsible for 25 percent of all enrollee costs above $30,000 and because they would have to present community-rated premiums during the annual enrollment period.

Rather than “premiums” being paid out of tax deductions or credits as in many proposals, a new federal agency would be created to administer premium collection and manage the pool in general. Premiums would be paid to this agency by enrollees. The additional federal revenue would not be scored as increased taxation, since there is no individual mandate. To minimize the chances of adverse selection in the absence of an
individual mandate, it would be essential to provide incentives for people to enroll when they are healthy rather than just when they are sick. These incentives would come in the form of a biennial enrollment period. Citizens who failed to enroll during this period would be able to enroll thereafter only if they experienced one of several life events (including loss of a job in which they had employer coverage) or if they paid a penalty in the form of higher premium (for which federal subsidies would be unavailable).

The federal government would subsidize on a sliding scale the premiums of families under 200 percent of the federal poverty line, with incentives for them to join managed care plans to promote cost control. Subsidies, then, (aside from some relatively small level of cross-subsidy by those with small claims of those with medium-size claims) would be fairly transparent in this system – federal revenues would redistribute toward the poor through premium support and toward the very sick through the reinsurance.

My proposal would provide access to affordable coverage for all Americans, would begin to move health insurance into a portable, individual-directed system, would help American employers compete with international rivals, would give workers more choices in terms of health insurance options (including the option to put higher wages to other uses and go without insurance), and would minimize the chances that the currently insured would end up worse off. By bringing greater transparency to the responsibility for health care costs, the proposal would set up a competitive insurance market whereby incentives for cost-conscious care might be introduced in the form of lower-cost plans in exchange for less comprehensive benefits or more tightly managed care.
No expansion of coverage is possible without redistribution of some sort, and my proposal would include a clear and explicit form of redistribution, all or partly in the form of additional revenue from ending the preferential tax treatment of employer-provided health insurance and the shift to greater earnings that would result. It would be relatively unlikely to draw fire from employer groups, insurers, or provider groups and would be less vulnerable to scare tactics from the right (much of which might actually support the proposal). My plan would accomplish all of this while remaining true to a number of deeply American values, expanding options while retaining the aspects of our health system that Americans like.

Retirement savings benefits could be delivered in a similar way. The federal government would sponsor 401(k)- or IRA-style plans (IRAs and similar individual accounts would be rolled into the new savings accounts). Individuals could contribute pre-tax earnings to these plans, but would not have to. Employers could match their employees’ contributions to these plans, but would not have to. They could also continue to sponsor their own retirement plans and match employee contributions to those plans. The federal government would provide its own match for contributions made by individuals under 200 percent of the poverty line. By default, all income tax refunds would automatically be placed in an individual account unless the taxpayer opted otherwise. Also by default, 3 percent of earnings would be transferred by employers from each employee’s paycheck to an individual account if the employee did not participate in an employer-sponsored plan or opt out of the default option.

The government-sponsored plans would be convenient in that workers changing jobs could simply leave their savings in their account without transferring them to a new
employer’s account, and there would be hefty penalties for withdrawing funds prior to retirement age. The government-sponsored plans would also include default allocation changes as an employee neared retirement age, which the person could again opt out of, to protect individuals from riskier investments that could wipe out their savings in a downturn.

The government-sponsored plans would include government-sponsored annuities as well, that the savings would be used to purchase by default (with an opt-out option). This provision could help create a functioning annuities market that overcomes the moral hazard problem that exists today, whereby the market is overwhelmed by buyers whose savings are unlikely to last long without an annuity, driving out buyers whose savings are likely to last them longer. Government-sponsored accounts might be complemented by a long-term care insurance component that individuals could be incentivized to purchase, which could help improve long-term care insurance markets too.

Social Security would continue to exist, but over time would shrink to become more of a safety-net pension program (along with retaining its other functions, such as disability insurance). Payroll taxes would eventually decline. The tax breaks given to individuals to induce them to save would constitute costs from a budgetary perspective, but would appear as tax cuts to participants rather than as taxes. There would also be some cost associated with the federal match. However, the higher savings rates might be expected to increase economic growth, which would potentially offset the cost to the Treasury.

Other potential products that could be offered as part of a menu of voluntary government-sponsored citizen benefits might include long-term care insurance,
catastrophic care insurance, or income-loss insurance. The latter could cover
unemployment, divorce, death of a spouse, short-term disability, or even lowered
earnings after a job loss.

The late Daniel Patrick Moynihan once said that, “Everyone is entitled to his own
opinion, but not to his own facts.” Many on the political right will argue that the
evidence I have compiled argues against even current policies to protect against
economic risk, let alone new ones. But as I noted at the beginning of the chapter, it is
unnecessary to show that some problem is getting worse to believe that it is too big. I
have outlined an approach friendly to the market to help those who become – or worry
about becoming – its victim. Arguing against policy aimed at economic risk requires
something more than reassurance that the sky is not falling any faster than in the past.

In response to my argument, many on the left might prefer to ignore these results
out of fear that acknowledging them will weaken the case for more activist federal
policies to address risk. Such a reaction would be a mistake for at least four reasons.
First, a message of “gloom and doom” does not resonate with a relatively secure citizenry
and will be counterproductive to winning political converts. We do not live in East
Germany in 1990. If Americans do not feel that progressives speak to their real
economic needs, they will be seduced by conservatives and their promise to return their
tax dollars. Second, focusing on the relatively minor anxieties of the middle-class may
result in fewer resources to address the bigger economic problems of the truly
disadvantaged.
Third, to the extent that progressives succeed in convincing Americans that they should be anxious about economic risk, the reaction could be something different than what the left expects. Rather than being a force for solidarity—a sense that we are all in this together—economic anxiety can provoke the opposite reaction. A strong case can be made that efforts to promote equality have been most successful when economic times are good and Americans are feeling generous.66 If Americans think that they have somehow managed to avoid calamities that are befalling many others around the country, they may become less willing to relinquish what they have in the name of collective security.

Finally, if progressives convince workers that we are living in times of heightened economic risk, those workers may be less willing to demand more from their employers in the way of wages, benefits, and better working conditions. Then-Federal Reserve Chairman Alan Greenspan famously cited job insecurity as a primary reason for low inflation in the late 1990s. Former Labor Secretary (and American Prospect founder) Robert Reich concurred.67

The defining characteristic of a progressive is not disregarding inconvenient facts or ignoring unintended consequences. Policies that would restructure vast segments of the economy on the basis of misperceived economic – and political – conditions may do more harm than good, from the perspective of both economics and politics. Perhaps the absence of a great risk shift does not require any changes in the progressive agenda. But if that’s true, then we should abandon the risk-shift narrative and make our case on other grounds.

Appendix One:  

Previous Research on Earnings Instability and Volatility Trends

Research on economic mobility and volatility has experienced something of a boom in recent years, driven by growing concern that the structure of the U.S. political economy has changed fundamentally in a way that leaves workers and families at greater economic risk than in the past. Nevertheless, scholars have examined trends in short-term earnings fluctuations for nearly two decades. I organize my review of the literature according to the groupings laid out in the introduction. ¹

Research on Short-Term Relative Mobility

We can track relative mobility in earnings back to the 1930s thanks to invaluable research by Wojciech Kopczuk, Emmanuel Saez, and Jae Song. Analyzing restricted-use earnings data from the Social Security Administration, these researchers charted trends in the probability of remaining in the top two quintiles or the bottom two quintiles in consecutive years, and in the probability of rising from the bottom two quintiles to the top two quintiles or falling from the top to the bottom. ² The figures in their charts are


² Kopczuk, Saez, and Song (2007). Their sample includes all commerce and industry workers (70 percent of U.S. employees) between the ages of 18 and 70, excluding a small number of very-low-wage workers (and everyone with $0 in earnings) and all self-employment earnings.
publicly available, and they can be used in combination to consider other measures of mobility.³

Figure A.1 uses their estimates to show three trends—the percentage experiencing downward mobility from the top two quintiles from one year to the next, the percentage experiencing upward mobility from the bottom two quintiles, and the probability of experiencing either change. Non-directional mobility defined by this latter measure declined dramatically from the mid-1940s until the early 1960s, rose through the mid-1970s, and then continued declining, ending at a low for the post-Depression era in the early 2000s. Upward mobility from the bottom shows a broadly similar trend. Downward mobility from the top, on the other hand, shows a smaller decline from the mid-1940s until the mid-1960s, with little secular change thereafter. There is evidence of a countercyclical pattern from the mid-1950s onward, with downward mobility temporarily increasing during recessions and upward mobility temporarily decreasing. These patterns mostly cancel each other out when looking at non-directional mobility.

In a newer version of their paper, Kopczuk et al. report trends since 1978 in the likelihood of moving down from the top percentile of earnings.⁴ Downward mobility declined somewhat over the late 1970s and early 1980s, increased in the mid-1980s, then gradually declined through the early 1990s. It rose again over the rest of the decade before falling again after 2000. Over the entire period there was little change in downward mobility.

³ The complete set of figures for their charts and tables are available at http://www.columbia.edu/~wk2110/uncovering/.

⁴ Kopczuk, Saez, and Song (2009). The methodological details are similar to their 2007 paper, except that the age restriction is tighter, excluding workers under 25 and over 60.
Maury Gittleman and Mary Joyce found that non-directional mobility among men and women declined a bit from 1968 to 1970, as measured by one minus the share remaining in their quintile in consecutive years or one minus the share either remaining in their quintile or moving into a neighboring quintile.\textsuperscript{5} Mobility increased in the first half of the 1970s, and over the rest of the decade mobility was flat among men but continued to increase among women. Robert Moffitt and Peter Gottschalk, also considering movement out of an initial quintile, found that among 36-year-old white male heads in

\textsuperscript{5} Gittleman and Joyce (1995). These results are based on matched March CPS data, with the sample including workers age 25-59 in both years who were not self-employed. The structure of the Current Population Survey is such that every March, half the sample was also interviewed in the previous March survey. Gittleman and Joyce match households to their earlier data for each survey from 1968 to 1992. The measure of earnings is wage and salary income, and those without earnings are dropped, along with the top one percent of earners.
the Panel Study of Income Dynamics (PSID), mobility declined in the early 1970s, increased briefly, then continued declining in the late 1970s. Mobility over a five-year period declined in the second half of the 1970s.\(^6\)

Gary Fields and his colleagues considered several measures of relative mobility, once again using the PSID and looking at male earnings.\(^7\) They found that the average centile change over five years rose from 1970-75 to 1975-80, indicating increasing mobility. Similarly, the share of men moving more than five centiles increased between these two periods, as did mobility measured as 1 minus the centile correlation coefficient or as 1 minus the rank correlation coefficient. Finally, when the men were classified into earnings deciles in both years, mobility measured as the negative of the chi-squared statistic for the ten-by-ten transition matrix also increased from 1970-75 to 1975-80.

Mary Daly and Greg Duncan examined the number of years in an eleven-year period in which a man remained in his original labor income quintile. They found higher mobility in the 1980s than in the 1970s.\(^8\) According to Gittleman and Joyce, non-directional mobility among men and women declined a bit through the mid-1980s and then flattened. Richard Burkhauser and his colleagues found that the Spearman rank

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\(^6\) Moffitt and Gottschalk (1995). The authors use PSID data from 1970 to 1988 to look at the probability of moving into another wage quintile over a 1-year period and over a 5-year period. Their sample consists of white men age 20-59 with positive wages and hours of work who are not students. They use the disadvantaged “SEO” oversample and weight their results. They trim the top and bottom 1 percent of wages within age-year cells and residualize wages by regressing log wages on education categories, stratified by age and year.

\(^7\) Fields, Leary, and Ok (2000). The authors examine the labor income of men age 25-60 in the first year of each five-year period who were not students, retired, or self-employed. They also exclude those with nonpositive earnings in either year.

\(^8\) Daly and Duncan (1997). The authors looked at the labor income of men in the PSID, including the disadvantaged SEO sample. They exclude men under 25 years old or over 44 years old in the first year of each of their eleven-year periods, those who had nonpositive earnings or less than 250 hours of work in the first year, and those who were self-employed. The number of years in which a man remained in his original quintile is expressed as a fraction of the number of years he was not self-employed.
correlation coefficient for logged labor income between consecutive years in the PSID was constant between 1984 and 1988. Moshe Buchinsky and Jennifer Hunt found that mobility declined during the 1980s, among both men and women, which is also consistent with Kopczuk et al., as is their finding that upward mobility from the bottom quintile declined more than downward mobility from the top quintile. Moffitt and Gottschalk also found declines in wage mobility over consecutive years among male heads in the 1980s, but when they measured mobility using five-year windows they found little change. Daly and Duncan, too, found mobility declines among men, from the 1979-89 period to the 1985-95 period. Finally, the measures used by Fields and his colleagues also show declines during the 1980s, which for the most part continue into the early 1990s.

In sum, it appears that short-term relative mobility declined during the postwar period, through the early 1960s, with downward mobility declining through the mid-1960s. Relative mobility then began increasing, probably reversing the early ‘60s decline, though there are inconsistent findings at the end of the decade. The patterns of non-directional relative mobility during the 1970s are inconsistent, but with the exception of Moffitt and Gottschalk’s study, they generally find increases in mobility. Non-directional mobility declined during the 1980s and in the early 1990s before flattening

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9 Burkhauser et al. (1997). They include the SEO sample of the PSID and weight the data. The authors dropped all workers with $0 in labor income in all years or who have a transition to or from $0 in any year, and only heads and partners between the ages of 25 and 55 are included in the sample.

10 Buchinsky and Hunt (1999). The authors use the National Longitudinal Survey of Youth, 1979 Cohort to examine mobility in terms of hourly wages during the survey week and annual earnings in the previous year. Mobility is measured as one-year transition rates between quintiles. They examine all workers not enrolled in school, not in the military, and not self-employed. They trim earnings that are more than five times the maximum or less than one-fifth the minimum of a worker’s earnings in other years, as well as wages below $1. Nonpositive earnings are excluded. Results are similar for hourly wages and annual earnings.
Research on Intertemporal Earnings Associations

The introduction noted that intertemporal earnings associations are best thought of as incorporating both relative and absolute mobility. As nondirectional mobility measures, they would be expected to follow trends in nondirectional relative and absolute mobility more closely than separate trends in upward or downward mobility. Most of these studies rely on the PSID. Many of the estimates come from correlation matrices included simply as descriptive statistics in research modeling earnings dynamics. Most are confined to men.

11 Depending on how relative mobility is measured, downward and upward mobility may be constrained to mirror each other, in which case trends in nondirectional relative mobility would follow trends in directional relative mobility. For example, if upward mobility was defined as the probability of leaving the bottom half of the distribution and downward mobility the probability of leaving the top half of the distribution, and if the population of interest was fixed over time, no one could move up without someone moving down. Research generally does not measure directional mobility in such ways however. For example, upward mobility from the bottom quintile need not mirror downward mobility from the top quintile because there is also mobility into the top and bottom quintiles from the middle 60 percent of the distribution. Furthermore, the population does not remain fixed over time – new households can enter the sample and households can leave the sample. Therefore, individuals’ percentile rank can change even if their earnings do not, if other people experience earnings gains or declines.
Kopczuk et al. computed rank correlations for commerce and industry workers and for men in particular, using SSA data. They found that earnings mobility fell in the late 1930s, rose more strikingly in the first half of the 1940s, then fell just as sharply in the subsequent years. Mobility continued falling modestly over the 1950s before flattening out in the 1960s. It rose in the late 1960s but then fell beginning in the early 1980s (from the mid-1970s to the mid-1980s for men). It changed little from 1985 onward.

Michael Baker showed the correlation matrix for male heads’ earnings over several years of the PSID. Short-term mobility as measured by year-to-year correlations increased from 1969 through 1972. It fluctuated between 1972 and 1976 before declining through 1982. Overall, mobility changed very little from 1970 to 1980. Steven Haider’s PSID correlations for men show little trend from 1968 to 1972, an increase in mobility through 1976, and a decline from 1978 to 1981. Like Baker, he found little change over the 1970s as a whole. Moffitt and Gottschalk also examined male autocorrelations over the 1970s, which show different patterns depending on whether they are measured over one or five years but which both indicate declining mobility over the decade.

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13 Baker (1997). The author excludes the SEO sample and includes male heads 21-64 with positive labor income and hours. He dropped men whose hourly wage was more than $100 in 1967 dollars or whose hours were greater than 4,680. Earnings are first regressed on year dummies and a quadratic in potential experience (age minus education minus 5).

14 Haider (2001). Haider excludes the SEO sample of the PSID and focuses on white male heads age 25 to 60 with positive earnings. Individuals must be observed at least twice from 1968 to 1992, and he excludes outliers in terms of hourly wages and observations with major imputations. He regresses log wages on a quartic in experience and displays the correlation matrix covering his sample years.

15 Moffitt and Gottschalk (1995). The appendix to the paper includes a full correlation matrix by age and year. See note 6 for methodological details.
On the other hand, Gittleman and Joyce found that mobility among men and women increased during both the first and second halves of the 1970s according to one-year autocorrelations using the CPS.16 Similarly, Daly and Duncan examined trends in autocorrelations using the PSID and found that male earnings mobility was higher in 1981 than in 1971.17 Fields and his colleagues reported mobility increasing from the first half of the 1970s to the second half, using five-year correlations in male earnings.18 Gittleman and Joyce, in a second paper, fit a quadratic trend through one-year wage and salary correlation coefficients within sex/age/education cells and within race cells in matched March CPS data.19 Their regression coefficients indicated that mobility increased during the late 1960s and through the 1970s.

Turning to the 1980s, Haider reported little trend over the decade but slightly lower mobility in 1990 than in 1980. Similarly, Gittleman and Joyce (1996) estimated that mobility declined between 1980 and 1990, though the change was small. Gittleman and Joyce (1995) found a decline in mobility among men and women, concentrated between 1982 and 1986. Moffitt and Gottschalk found increases in mobility in the early 1980s followed by slightly bigger declines mid-decade. Fields and his colleagues also found declining mobility in the 1980s and through the early 1990s.

16 Gittleman and Joyce (1995). The correlations are presented without conditioning on other variables, individually for each year. For additional methodological details see note 5.

17 Daly and Duncan (1997). Autocorrelations are for earnings separated by two years. For additional methodological details, see note 8.

18 Fields, Leary, and Ok (2000). See note 7 for methodological details.

19 Gittleman and Joyce (1996). Gittleman and Joyce match households to their earlier data for each survey from 1968 to 1992 and compute correlation coefficients for those ages 25-54 in the first year. They exclude the self-employed and those who have nonpositive earnings in either year and trim the top 1% of wages within sex groups. In one set of analyses, correlation coefficients are conditional on sex, age, and education; in a second set they are conditional on race. These correlations are regressed on a number of variables, among them a quadratic in time.
On the other hand, Baker found that male heads’ mobility increased from 1982 to 1986 after several years of decline. Daly and Duncan’s PSID autocorrelations also indicate an increase in mobility over the mid-1980s. In yet another study using the PSID, Dean Hyslop reported autocorrelations of husbands’ and wives’ hourly wages and earnings.\(^{20}\) Wage and earnings mobility among husbands were higher in 1985 than in 1980. Among women, wage mobility changed little, while earnings mobility declined.

Finally, Bhashkar Mazumder used a unique dataset that consists of Survey of Income and Program Participation (SIPP) data matched to Social Security Administration data on earnings. He reported that mobility as measured by year-to-year autocorrelations of logged residualized earnings among men declined from 1984 to 1990, increased from 1990 to 1992, and then declined from 1992 to 1997.\(^{21}\)

The various estimates of intertemporal earnings association are remarkably inconsistent, particularly given that so many of them rely on the PSID and are focused on men. Baker, Haider, and Kopczuk et al. agree with one another that mobility increased in the late 1960s and declined in the 1970s, but they are contradicted by Gittleman and Joyce on both counts. In regard to the 1970s, Moffitt and Gottschalk side with Baker, Haider, and Kopczuk et al., but Daly and Duncan side with Gittleman and Joyce. In the 1980s, Haider’s and Baker’s trends conflict, as do those of Daly and Duncan and of Gittleman and Joyce, but Haider’s agree with Gittleman and Joyce, as do those of

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\(^{20}\) Hyslop (2001). Hyslop included the SEO. The sample includes husbands and wives age 18-60 in 1980 who were continuously married over the period, with no top-coded labor income data and no years with wages higher than $100 an hour. Both husbands and wives must have positive labor income and hours in every year. Hourly labor income and labor income were adjusted for inflation using the CPI and year-specific means were removed from individual earnings before computing correlations.

\(^{21}\) Mazumder (2001). The author pooled the 1984, 1990, and 1996 waves of the matched SIPP data. His sample includes men who were 24-59 for at least nine consecutive years between 1982 and 1998 in which they also had positive earnings, but he drops the first and last years of earnings data, so sample members are 25-58.
Kopczuk et al. Mobility likely declined in the second half of the 1980s. Finally, Mazumder’s results imply that mobility increased in the early 1990s before declining through the middle of the decade, ending lower in 1997 than in 1990. Kopczuk et al., on the other hand, find little change over the 1990s. There is too much uncertainty in these estimates to compare them against the estimates of non-directional relative mobility. However, Kopczuk et al.’s estimates are consistent with their relative mobility results indicating major declines in mobility prior to 1960.

Research on Short-Term Absolute Mobility

An alternative to looking at short-term relative mobility trends or trends in mobility as intertemporal association is to examine trends in absolute mobility. Beginning with downward absolute mobility, Molly Dahl, Thomas DeLeire, and Jonathan Schwabish analyzed the restricted-use Social Security Administration (SSA) data to examine trends in the probability of experiencing large earnings drops going back to the early 1960s, though due to data limitations, the analyses were restricted to the bottom 40 percent of the wage distribution.22 They found that for both men and women, downward wage mobility (experiencing a decline of either 25% or more or 50% or more over the course of a year) declined from 1961 to 1966 and increased from 1966 to 1971, consistent with trends in relative downward mobility from the SSA data. Absolute downward mobility was slightly higher in 1971 than in 1961, while relative downward mobility was significantly higher. The cyclical swings Dahl and her colleagues found in

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22 Dahl, DeLeire, and Schwabish (2007). The authors looked at workers age 22 to 59, using the Continuous Work History Sample of the Social Security Administration. Self-employment earnings and deferred compensation are not included, and workers with $0 in earnings in both years are also excluded. If a worker had $0 in the first year but not in the second, the increase was coded as 100 percent.
the 1970s among men and women also mirror the downward relative mobility trends for all workers. The similarity in downward mobility rates in 1970 and 1980 masks a sizable increase among men, with a smaller decline among women.

The work of Karen Dynan and her colleagues, using the PSID, shows that downward labor income mobility among family heads increased in the mid-1970s before falling again.23 Jacob Hacker has also produced estimates of the likelihood of large drops in labor income using the PSID, and they also show downward mobility rising and falling cyclically in the 1970s.24 Dynan et al. and Hacker each show a small increase over the decade.

In the 1980s, as with relative mobility trends based on the SSA data, Dahl et al. found that absolute downward mobility rose and fell early in the decade and then was relatively flat (slightly increasing), ending the decade lower than where it began. This was true for both men and women. From 1980 onward, the SSA data may be used to look at all workers, rather than having to trim the top three quintiles. Dahl et al. showed that downward mobility increased from 1981 to 1982 (as in the trimmed data), then declined through 1985 among both men and women. The full SSA data shows a slight decline from 1986 to 1990, basically consistent with the trimmed results.

23 Dynan, Elmendorf, and Sichel (2007). Mobility is measured as the share of heads experiencing an annual drop in labor income of 25% or more per year over a two-year period. In computing percent drops, the change in earnings between years $t$ and $t-2$ is divided by the average of years $t-2$, $t-3$, and $t-4$. The sample includes individuals who were heads in both years, at least 25 years old and not retired, and individuals with farm income are excluded. Incomes are bottom-coded at $1$, and top codes are applied to cap the same share of the sample in each year. In all numbers cited in this appendix, I refer to results when reports of $0$ in earnings by the head are dropped. Dynan et al. do not include the disadvantaged “SEO sample” or the immigrant sample of the PSID.

24 Hacker (2007a). Hacker includes the SEO and immigrant samples, logs labor income, applies top codes to cap a common share of incomes in each year, and excludes those with nonpositive labor incomes. His mobility measure is the probability of experiencing a 50% drop in labor income over a two-year period. Individuals between the ages of 25 and 61 are included.
Dynan et al. also found continuing cyclical patterns in the first part of the 1980s, with downward mobility rising in the early 1980s before falling yet again. They found that downward mobility in the PSID was flat during the late 1980s. Hacker, too, found cyclical patterns through the early 1980s, and his trends for the rest of the decade are consistent with the other studies (and similar for men and women). Like Dynan et al. and Hacker, but in contrast to Dahl et al., Daly and Duncan’s PSID results show higher downward mobility in the 1980s than in the 1970s, measured as the average across male workers of the number of years that a person experienced a 50-percent earnings drop in an 11-year period.25

Dahl et al.’s absolute mobility trends show a longer and stronger decline in the 1990s than the SSA-based relative mobility trends. Like the trimmed SSA data, the full data shows a small increase in downward mobility in 1991 before declining through the late 1990s. Finally, both samples show small increases in downward mobility from 2000 to 2003. As with the relative mobility trend, however, the downward absolute mobility trend ended near early-1990s levels.

Dynan et al. found that downward mobility in the PSID rose in the early and mid-1990s. It then declined for the next few years, increasing at the end of the decade. They show increasing downward mobility through 2002, with a small decline from 2002 to 2004. Again, the SSA and PSID patterns are cyclical and consistent, except that the PSID shows an upward secular trend after 1990. While the SSA results show downward mobility in the early 2000s to be lower than its early 1980s levels and comparable to mid-1970s levels, Dynan et al. show it to be higher than any previous period.

Hacker’s trends for the early 1990s are unique. He shows larger downward mobility increases among both men and women than in the other studies. Downward mobility then declined over the rest of the decade, increasing after 1998 among men. Hacker shows an increase from 2000 to 2002 among men and women, followed by an equally large decline among men from 2002 to 2004 (no change among women). Combining men and women, his results show a big increase in downward mobility followed by a decrease that is only a little less dramatic than the preceding increase.

Hacker has never presented trends in the likelihood of large earnings gains, but Dahl et al. and Dynan et al. have. Dahl and her colleagues found that from 1981 to 1987, the probability of a 25-percent or 50-percent increase in earnings decreased slightly for women and increased slightly for men (converging), but then it declined notably through 1991 for both. It then increased more shallowly through 1998 and decreased through 2002, ending up slightly lower than in 1991 for women and slightly higher for men. All of these patterns are similar to those for relative upward mobility in the SSA data.

Dynan et al., examining the likelihood of a 44% increase in earnings over two years among family heads (25% per year), confirmed a cyclical pattern from the 1970s forward. They found little change in upward mobility in the 1970s. Contrary to the SSA-based relative and absolute upward mobility results, Dynan and her colleagues found little decline in the 1980s, an increase in the 1990s where the SSA data shows none, and a big drop post-2000.

Two studies provide estimates of the probability of experiencing a large earnings change in either direction (up or down). Dahl et al.’s results imply a decline in the 1980s for men and women, with a flat trend in the 1990s through the early 2000s among men
and a small decline among women. Once again, these trends conform with the trends in non-directional relative mobility. In contrast, the results of Dynan et al. indicate that the percent of family heads experiencing a large change in earnings increased steadily from the early 1970s through the 1990s, then declined somewhat.

In addition to these upward and downward absolute mobility studies, other research focuses on non-directional measures of absolute mobility. A recent paper by Shane Jensen and Stephen Shore measured “volatility” as the squared change (over two years) in logged residualized labor income, using the PSID.\(^\text{26}\) It found that mean and median volatility among male heads increased from 1969 to 1975. Mean volatility showed little trend between 1975 and 1991 but then jumped through 1994 before falling through 2000. Median volatility declined from 1975 to 1976 but then showed little consistent trend through 2000. Both mean and median volatility jumped in 2002. The difference between the mean and median trends was driven by changes in large income swings – the trend in the 95\(^{\text{th}}\) percentile of labor income volatility closely followed the trend in mean volatility.

Fields and his colleagues used the PSID to look at the mean of the absolute values of earnings changes.\(^\text{27}\) Whether they measured earnings in terms of levels or logs, mobility increased during the 1970s and through the first half of the 1980s. Mobility in terms of levels was flat over the 1980s and then declined from the late 1980s through the early 1990s. Mobility in terms of logs declined from the early 1980s to the early 1990s.

\(^{26}\) Jensen and Shore (2008). The authors include the SEO and applying survey weights. They focus on male heads age 22-60. Labor income is bottom-coded at $5,150 in 2005 dollars (1,000 hours at the 2005 minimum wage), and a common topcode is applied across all years (the minimum real topcode level across years). The log of labor income is residualized by regressing it on a cubic in age interacted with 8 educational attainment levels, a dummy for the presence infants, young children, and older children, the number of children in each group, the number of family members, and year dummies.

\(^{27}\) Fields, Leary, and Ok (2000). See note 7 for methodological details.
It is difficult to generalize across these absolute mobility findings, but there is little evidence that earnings instability has recently become a new structural problem. Dahl et al. find that downward mobility declined during the first half of the 1960s and then increased, declining over the decade as a whole. Non-directional absolute mobility increased in the 1970s, apparently due to an increase in downward mobility among men. However, the 1980s featured little change in upward or downward mobility, or perhaps declines in one or both. The non-directional measures of absolute mobility are inconsistent, but only two out of six show increasing instability over the decade. Results from the 1990s are also inconsistent, but it appears that there was either little change or a decline in downward mobility at the same time that there was little change or an increase in upward mobility. The first years of the current decade, marked by the 2001 recession, featured increases in downward mobility and declines in upward mobility. As with directional measures of relative mobility, downward and upward absolute mobility show countercyclical patterns.

**Research Examining the Dispersion of Earnings Changes**

The studies in this line of research all focus on the variance or standard deviation of changes in earnings. In a paper devoted to estimating a formal model of earnings change, John Abowd and David Card presented variances of men’s labor income changes for a number of years in the 1960s and 1970s, using two different surveys.\(^28\) They found

\(^{28}\) Abowd and Card (1989). The authors use the PSID and National Longitudinal Survey of Men 45-59. In the PSID, they show results including and excluding the SEO. The sample consists of males who were heads age 21-64 and reporting positive labor income and hours in every year from 1969 to 1979. It excludes workers with hourly wages over $100 an hour or hours greater than 4,680. In the NLS, the
that dispersion in one-year earnings changes in the National Longitudinal Survey of Men 45-59 increased from 1967 to 1975 by about 75 percent when the variances are expressed as standard deviations. In the PSID, dispersion of earnings changes declined between 1970 and 1974, increased through 1976, and declined again through 1979, ending about where it started.

Baker also presented PSID trends in the variance of male heads’ labor income changes between consecutive years. He reported that dispersion of changes increased from 1968 to 1972. It declined through 1974 and then increased through 1976 before declining again through 1980. Earnings-change dispersion increased from 1980 to 1986, ending at its highest level. Overall, when expressed as standard deviations, dispersion doubled between the late 1960s and 1986 (but increased by just 50 percent between the late 1960s and 1985).

Stephen Cameron and Joseph Tracy pursued a similar approach using interviews from March Current Population Survey respondents that were one year apart. They examined trends in the variance of changes in men’s logged wage and salary income after adjusting for age, education, and industry. Cameron and Tracy captured residuals from a model of the level of earnings in each year and then differenced the residuals from

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30 Cameron and Tracy (1998). The authors focus on men who are initially 18-63 and who are not in school nor primarily self-employed in either year. Workers with nonpositive weeks worked, nonpositive earnings, or earning imputations are excluded. The top and bottom 1.5% are trimmed in each year. The measure of volatility is computed by first regressing logged real wages and salaries on a quartic in age, four education groups, and six industry groups. The residuals are estimates of unobserved permanent income plus transitory income. The authors then subtract residualized wages in year \( t-1 \) from residualized wages in year \( t \), which gives the change in the transitory component. They square the difference, then regress the squared differences on year dummies to get the mean squared difference for each year.
respondents’ two interviews. They found that dispersion of earnings changes increased from 1967 to 1982, with cyclical spikes in the early 1970s and mid-1970s and an especially large increase from 1981 to 1982. From 1982 to 1996, dispersion declined somewhat, with a cyclical spike in the early 1990s. Expressed in standard deviations, dispersion increased about 20 percent from 1967 to 1996. The increase they found over the 1980s is at odds with the decline found in the SSA data.

Dynan et al. (2008) also relied on a similar measure.31 When heads and spouses were pooled, the standard deviation of percent changes in earnings in the PSID declined by 3 percent between the early 1970s and the early 2000s. However, while it declined by 15 percent among women (including both heads and spouses), dispersion of changes actually rose 70 percent among men. They found that dispersion among family heads increased steadily from the early 1970s to the early 2000s, by about 40 percent, though they noted that the increase since 1980 was smaller if they excluded those with a business interest, and their 2007 draft indicated that the increase fell to about 20 percent if those with a business interest were excluded, implying that self-employment earnings are more unstable than wages and salaries.32 The increase in earnings-change dispersion was greater before 1985 than after, and the 2007 draft indicated it was confined to male heads.

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31 Dynan et al. (2008). Methodological details are the same as for Dynan et al., 2007 with two exceptions. Rather than excluding heads reporting $0 in earnings, they exclude heads who report no earnings but who report over 120 hours of work in the previous year. The authors measure percent changes as the two-year change in earnings, divided by the average of the two years of earnings.

32 Dynan et al. (2007). The authors measured percent changes as the two-year change in earnings, divided by the average of earnings 2, 3, and 4 years earlier. Increases were capped at 100 percent. For the sample pooling heads and spouses, in addition to dropping heads with $0 reports, they drop spouses with two successive years of $0 reports. For additional methodological details, see footnote 23. Dynan et al. (2008) also reported that volatility in family heads’ earnings per hour increased between the early 1970s and the early 2000s, with the increase almost exclusively confined to the mid- to late-1970s and the early 1990s.
Dispersion among the spouses of family heads declined by about 15 percent over the period, spouses in the PSID being overwhelmingly women.

In another paper in this line of research, Donggyun Shin and Gary Solon computed the standard deviation of the two-year change in log earnings for male heads, after adjusting earnings for age.\(^{33}\) They found that volatility increased in the 1970s but found little secular trend from the late 1970s to the late 1990s. Volatility increased from 1980 to 1983, then declined through the mid-1990s. It increased substantially after 1998, ending over 50 percent higher in 2004 than in 1971.

Moffitt and Gottschalk presented trends in the variance of two-year residualized earnings changes among male heads age 33 to 35, again using the PSID.\(^{34}\) Dispersion increased in the first part of the 1970s, declining in the latter part of the decade. It then increased in the early 1980s before declining over the rest of the decade (with a big spike in 1986). This cycle repeated in the 1990s, with an increase early in the decade followed by a decline through 1998. Earnings-change dispersion rose from 1998 to 2002 and declined from 2002 to 2004. Over the entire period, the increase from 1973 to 2004 was about 40 percent (expressed in standard deviations), again following a countercyclical pattern.

\(^{33}\) Shin and Solon (2009). The authors regress the two-year change in log wages and salary on age and age squared, separately for each year. They then compute the mean square of the residuals for each year and use the square root of that to report results in standard deviations. They drop all observations with $0 in earnings and trim the top and bottom 1 percent of positive observations. Shin and Solon look at wage and salary income of male heads age 25-59 in both years. They exclude the SEO sample.

\(^{34}\) Moffitt and Gottschalk (2008). The authors rely on the PSID, excluding the SEO sample. The sample consists of male heads age 30-59 who are not students, who have positive wage and salary income and who have positive weeks worked. They look at log wage and salary income adjusted for inflation using the CPI-U-RS and residualized on year, education, race, a polynomial in age, and interactions among these variables. They trim the top and bottom one percent of residuals within age/education/year cells.
Moffitt and Gottschalk also examined the trend in the variance of ten-year earnings changes for 35- to 45-year-olds. The variance declined during the early 1980s, increased through the mid-1980s, and then declined over the rest of the decade. It then increased through the mid-1990s and declined through 2000. It increased from 2000 to 2002 and declined from 2002 to 2004. The increase from 1980 to 2004 was about 20 percent (versus no change when two-year changes were examined).

Another paper using the PSID also focuses on dispersion in residualized earnings. Susan Dynarski and Jonathan Gruber adjusted male heads’ earnings for a number of demographic variables before looking at trends in the variance of the one-year change in male heads’ logged earnings.\(^{35}\) Expressing their results in standard deviations rather than the variances they present, dispersion rose about one-third from 1970 to 1991, though the trend clearly follows a countercyclical pattern.

Finally, Dahl et al. included time series for men and women using a measure based on dispersion in earnings changes.\(^{36}\) Using the SSA data again, they found that the standard deviation of one-year percent changes in earnings declined 10 to 12 percent from 1981 to 1991 among both men and women. Updated figures show a roughly 10-percent decline from 1985 to 1991 among both groups.\(^{37}\) The standard deviation of the difference in log earnings declined 8 percent among men and 18 percent among women.

\(^{35}\) Dynarski and Gruber (1997). The authors regress the one-year change in log labor income on year dummies, a quartic in age, three education groups, marital status, change in marital status, change in family size and in the share under 18, and change in food needs. They then use the mean squared residuals for each year. Their sample includes male heads ages 20-59 who are not fulltime students, and they use the SEO sample and weight their results.

\(^{36}\) Dahl, DeLeire, and Schwabish (2007). In computing percent changes, workers with no earnings in both years are excluded, and if one goes from no earnings to positive earnings, it is coded as a 100% increase. Workers with changes in earnings of 1000% or more are excluded. See footnote 22 for additional methodological details.

\(^{37}\) Dahl, Schwabish, and DeLeire (2008). These updated results examine workers age 25-55 and use the average of years \(t\) and \(t-1\) for the denominator in computing percent changes.
from 1981 to 1991 (5 percent or less from 1985 to 1991). Dispersion of changes then increased slightly over the next few years, only to decline slightly over the rest of the decade, ending lower than in 1990. It increased a bit among men between 2000 and 2003 and was flat among women. Dispersion was about 3 or 4 percent lower in 2003 than in 1985, for both men and women. Much of the pre-1991 decline in earnings-change dispersion may be due to a parallel decline in the number of workers who have no earnings in the year preceding or following a year in which they worked.38

In sum, Moffitt and Gottschalk, Dynan et al., Shin and Solon, Dynarski and Gruber, and Cameron and Tracy all find increases in male earnings movement during the 1970s, which is consistent with the research on relative and absolute downward mobility showing an increase among men. Abowd and Card as well as Baker indicate little change in earnings-change dispersion among men. Dynan et al. find a small decline for women, consistent with the CBO results for downward mobility of women.

Moffitt and Gottschalk (looking at two-year changes), Shin and Solon, Dynarski and Gruber, and Cameron and Tracy all find a big increase in male earnings movement during the early 1980s recession, followed by declines over the mid-1980s. Dynan et al. find a similar pattern among male heads and spouses, though their earlier results indicated a fairly steady increase in earnings-change dispersion among male heads over the first half of the 1980s and a leveling off during the last years of the decade. Shin and Solon show a deeper decline during the middle part of the decade than the other studies. According to Dynan et al.’s results, earnings movement among women declined steadily during the 1980s. These results for men and women are in some sense consistent with

38 Orszag (2008).
the SSA-based estimates, but the early-1980s increase in earnings movement is swamped by the decline over the rest of the decade in the SSA data, while the PSID studies—and Cameron and Tracy—find an increase over the 1980s among men (except for Moffitt and Gottschalk). Dahl et al. find fairly steady declines among both men and women over the decade, consistent with their own findings on absolute mobility.

Shin and Solon and Dahl et al. find a sizable decline in earnings movement during most of the 1990s, with a small and temporary increase in the early 1990s, while Cameron and Tracy also show the increase. Moffitt and Gottschalk also found an increase followed by a decline, but they showed earnings-change dispersion rising over the decade. Dynan et al. find a bigger increase in earnings-change dispersion among heads and men in the early 1990s, followed by a smaller decline. They find declines among women, as in the Dahl et al. study.

Finally, Dahl et al. show a small increase in earnings movement among men after 2000, while Dynan et al. also find an increase among men. Neither study shows an increase for women. Shin and Solon report a large increase in earnings-change dispersion among male heads sometime after 1998. Moffitt and Gottschalk find a decline from 2000 to 2004.

Taken as a whole, the research on dispersion in earnings changes, contrary to the research on mobility, indicates that earnings changes continued increasing through the early 1990s, at least if one ignores the SSA-based research. It is unclear whether this increase is due to volatility or downward mobility becoming more severe or to a general increase in inequality or earnings growth, which would produce the same results. Different factors may have operated in different periods. On the other hand, the SSA-
based research may be the most valid, in which case, dispersion in earnings changes has not increased over the past twenty years. The research generally agrees that earnings-change dispersion, as with directional mobility, follows a countercyclical pattern, increasing during recessions and declining during expansions.

**Research Summarizing Within-Person Earnings Dispersion across Years**

The essence of the concept of volatility involves fluctuation within a short period of time. Various researchers have measured volatility as the typical dispersion of earnings individuals experience within some brief window of time. Nearly all of the studies in this line of research use the PSID. Peter Gottschalk and Robert Moffitt were the first to use the approach, and they introduced two measures that were subsequently adopted by others. In one analysis, Gottschalk and Moffitt estimated trends in the average white male head’s logged wage and salary income variance over a nine-year window. They found that from the nine-year period 1970-1978 to the nine-year period 1979-1987, volatility increased 42 percent (which would have been smaller if they had expressed the results as average standard deviations).

The second set of estimates was similar, except that they were based on a five-year window centered on a given year that moved over time. In any year, an individual’s transitory variance was the variance of his earnings across the current year, the preceding two years, and the subsequent two years. This approach is more flexible in that one can construct a time series rather than comparing two discrete periods. Gottschalk and

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39 Gottschalk and Moffitt (1994). The authors include the SEO sample of the PSID and weight the data. Their sample is restricted to white male heads not in school with positive wages and between the ages of 20 and 59. They trim the top and bottom 1 percent of wages within age/education/year cells. Prior to computing variances, they regress log wages on a quartic in age and individual fixed effects, and they use the age-residualized estimates in all computations.
Moffitt reported that volatility increased 120 percent between 1974 and 1986 but provided no detail on the evolution of the trend. Recently, however, they updated these results.\textsuperscript{40} They found that earnings volatility rose from the early 1970s to 1989, declined through 1998, and then rose again between 1998 and 2000. There was very little change between 1990 and 2000. Over the entire period, volatility rose about 85 percent.

Gottschalk and Moffitt earlier presented similar results using the Social Security Administration’s Continuous Work History Sample, estimated with the help of the Congressional Budget Office’s Jonathan Schwabish.\textsuperscript{41} Defining the window as spanning years $t-7$ to $t$, they found volatility declining from 1987 through the early 1990s, then flattening out before declining again in the late 1990s and early 2000s. The corresponding results using the PSID showed flat volatility from 1987 to the early 2000s, but rising volatility from the late 1970s to the mid-1980s.

Daly and Duncan adopted Gottschalk and Moffitt’s first approach to examine volatility changes among men, defining volatility as the average within-person variance of deviations from permanent earnings measured over an eleven-year period.\textsuperscript{42} Like Gottschalk and Moffitt, they found that volatility was higher in the 1980s than the 1970s.

Diego Comin and his colleagues also used Gottschalk and Moffitt’s first approach to examine volatility changes among white male workers in general and among those that

\textsuperscript{40} Moffitt and Gottschalk (2008). Volatility is measured as the mean within-person variance of earnings over a nine-year moving window. For additional methodological details, see note 34.
\textsuperscript{41} Gottschalk and Moffitt (2007).
\textsuperscript{42} Daly and Duncan (1997). For methodological details, see note 8.
were continuously employed by the same firm.\textsuperscript{43} They look at within-person dispersion of heads’ ten-year hourly wages. Among male workers in general, volatility increased 67 to 72 percent between the 1970-79 period and the 1984-93 period. What is more, volatility increased 79 to 115 percent among those who did not change jobs, implying either that the measure is picking up mostly higher wage growth in the latter period or that volatility is increasingly significant even among steadily employed workers (due to greater fluctuation in hours, for instance).

Comin and his colleagues also looked at five-year windows instead of ten-year ones. They found that white male heads’ earnings volatility increased 28 percent from the 1970-74 period to the 1975-79 period, increased 28 percent again from 1975-79 to 1980-84, declined by 1 percent from 1980-84 to 1985-89, and increased 20 percent from 1985-89 to 1990-94 (growing by 93 percent over the entire period). The changes were bigger for those staying in their jobs.

Comin et al. also adopted Gottschalk’s and Moffitt’s second approach, using overlapping five-year windows. In their most recent version of their paper, they extend these time series and also look at overlapping ten-year windows.\textsuperscript{44} Using the five-year windows, they find that volatility increased over the 1970s and through the early 1980s. It then declined before increasing again in the late 1980s and early 1990s. Volatility then declined over the mid- to late 1990s. Using ten-year windows, volatility increased from the mid-1970s to the early 1980s, was flat over the mid-1980s, rose again in the late

\textsuperscript{43} Comin et al. (2006). The authors include the SEO sample of the PSID and weight the data. They restrict their sample to heads and look at logged wages (earnings/hours). They adjust for inflation using the PCE deflator.

\textsuperscript{44} Comin et al. (Forthcoming). The authors include the SEO sample of the PSID and weight the data. They restrict their sample to heads and look at logged wages (earnings/hours). They adjust for inflation using the PCE deflator. The results I cite use every other year within a window, so that the five-year variances use years $t-2$, $t$, and $t+2$, and the ten-year variances use years $t-4$, $t-2$, $t$, $t+2$, and $t+4$. 
1980s and early 1990s, then declined through the mid-1990s. The authors find essentially the same trends when women and nonwhites are included in their samples and when they restrict the sample to people who stayed in their jobs. Among all individuals, volatility increased 40 to 70 percent over the entire period.

Benjamin Keys also recently updated Gottschalk and Moffitt’s findings based on their first approach, comparing transitory earnings in the 1970s, 1980s, and 1990s by looking at white male heads’ within-person variance over each decade. He found a 43 percent increase in volatility from the 1970s to the 1980s, a 4 percent decrease from the 1980s to the 1990s, and an overall increase of 38 percent from the 1970s to the 1990s. Volatility increased much more among black male heads (90 percent between the 1970s and the 1990s), and less among female heads (9 percent among whites and 17 percent among blacks). Among black women, volatility was flat between the 1970s and 1980s and only then increased between the 1980s and 1990s.

Peter Gosselin and Seth Zimmerman computed the variance of individuals’ earnings across four years in a seven-year window and then tracked changes in the average variance across individuals. They found a ten percent increase in volatility from 1970 to 1998, concentrated in the second half of the 1980s, with a decline in the 1990s. There was little change over the entire period. The authors also tracked changes in volatility using the SIPP, computing individual variances over three-year windows.

45 Keys (2008). Keys includes the SEO sample of the PSID and weights the data. He restricts his sample to non-student heads age 20-59, regresses log earnings on a quartic in age and uses the residuals, and trims the top and bottom 1 percent of positive earnings.

46 Gosselin and Zimmerman (2008). The authors include the SEO sample of the PSID and weight the data. They restrict the sample to adults between the ages of 25 and 64 and exclude those in households reporting $10 or less of income before out-transfers. Before estimating volatility, they regress log earnings on a quartic in age and use the residuals. Their within-person variance measure uses incomes in the current year and incomes two, four, and six years into the future.
They found an increase of 10 percent from 1983 to 2001. If one compares years for which Gosselin and Zimmerman have estimates in both datasets, the PSID shows an increase of 22 percent between 1983 and 1996, while the SIPP shows an eight-percent increase.

Austin Nichols and Zimmerman, once again using the PSID, measured volatility as the within-person standard deviation of three years’ of earnings.47 They found that, using earnings levels rather than logs, the average standard deviation increased about 50 percent between the early 1970s and the early 2000s. The increase accelerated over most of the period, with the 1990s showing the largest increase and a decline in the early 2000s. The median standard deviation, however, increased just 15 to 20 percent over the thirty-year period. When the authors examined trends in logged earnings dispersion rather than trends using levels, they found that the median was flat over the period, while the mean increased just 10 percent. Results using other transformations of earnings (the neglog, inverse hyperbolic sine, and cuberoot) were consistent with the logged results.

A working paper by Nichols and Melissa Favreault uses the SIPP matched to Social Security Administration earnings data to estimate the mean within-person variance over five-year periods.48 Nichols and Favreault find that mean volatility declined or was flat for five of six SIPP panels examined between 1977-81 and 1999-2003, though the trends across the panels defy any attempt to characterize the “true” trend for specific

47 Nichols and Zimmerman (2008). The authors include the SEO sample and weight their results, and they include individuals age 25-61 years old. Their measures consider the mean and median of the within-person standard deviation in earnings over three years (t-2, t, and t+2). Sample members must have non-missing data in all three years.

48 Nichols and Favreault (2008). The authors use the 1990-93, 1996, and 2004 SIPP panels, matched to the Summary Earnings Record and Detailed Earnings Record of the SSA data. The sample includes adults born 1936-1956 (21-41 years old in 1977, 48-68 in 2004). The largest one percent of variances are dropped. I thank the authors for providing me with the latest draft of their paper.
periods. Most panels showed large declines in volatility from 1977-81 to 1980-84 and increases in volatility between 1980-84 and 1999-2003. At the 75th percentile, volatility declined from 1977-81 to 1990-94 and then increased through 1997-2001, ending higher than it started.

Gosselin includes one other PSID analysis of earnings volatility, using a novel measure, in his recent book, *High Wire: The Precarious Financial Lives of American Families*. Gosselin examines trends in the maximum absolute value of annual percent changes over a five-year window. He reports the 68th percentile, separately for men and women, of this maximum. Gosselin found that volatility increased among men from the early 1970s through the mid-1990s, accelerating in the early 1990s, before declining through 2002. From the early 1970s to the early 2000s, the 68th percentile of the maximum swing increased from about 26 percent to about 35 percent. Among women, volatility declined through the late 1980s, then mirrored the trend among men, finishing at 40 percent (versus about 48 percent in the early 1970s).

The studies in this line of research that follow Gottschalk and Moffitt’s first approach are a bit difficult to compare to other studies, but the general finding that male volatility was higher in the 1980s than in the 1970s is consistent with other research. Comin et al.’s finding that volatility was lower in the second half of the 1980s than in the first half and Keys’s finding that male volatility was lower in the 1990s than the 1980s are also consistent with past research, with the sole exception of Dynan et al. (2008).

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The results of Gottschalk and Moffitt and Comin et al., using Gottschalk and Moffitt’s second approach, align with other studies in showing increasing volatility among men in the 1970s, and Comin et al. find the same early-1980s and early-1990s countercyclical increases. Both sets of authors’ PSID-based findings of an increase in volatility in the 1980s aligns with PSID-based studies that look at the dispersion of earnings changes. Gosselin also shows rising volatility among men in the 1970s and 1980s. He finds a small increase over the 1990s and early 2000s, while Gottschalk and Moffitt find little change over the 1990s (or declines using the SSA data). Comin et al. find flat or declining volatility among white men over the 1990s. Finally Gosselin finds that women’s earnings volatility changed little in the 1970s but declined in the following decades—generally consistent with the studies that examine dispersion in earnings changes.

The remaining studies combine men and women. Comin et al., Nichols and Zimmerman, and Gosselin and Zimmerman find increases in volatility in the 1970s. The 1980s estimates are all over the map. Comin et al. found an increase driven by cyclical patterns. Gosselin and Zimmerman found a flat trend followed by a late-1980s increase using the PSID but a slight decline over most of the decade using the SIPP. Nichols and Favreault found a decline over the decade. Nichols and Zimmerman found different trends depending on whether they looked at averages or medians and whether they considered earnings levels or logs. Nichols and Zimmerman, as well as Nichols and Favreault, find increases in volatility in the 1990s, but Gosselin and Zimmerman find a decline using the PSID, and Comin et al. report flat or declining volatility. Nichols and Zimmerman and Nichols and Favreault find declines in the early 2000s. All in all, while
volatility measured as within-person dispersion seems to have increased in the 1970s and perhaps in the 1980s among men, it is unclear whether it was higher in the early 2000s than in 1980, though there is evidence that it was among men, driven by increases through the 1980s or early 1990s.

**Research Summarizing Across-Person Dispersion of Earnings Shocks**

As noted in the introduction, a sizable literature models earnings dynamics as processes subject to “shocks”, and dispersion in these shocks constitutes another measure of volatility. Nearly all of the research in this tradition focuses on male heads, and nearly all of it relies on the PSID. Once again, the innovators of this line of research were Gottschalk and Moffitt. Beginning with their 1994 paper, they developed simple models estimating transitory variances that were forerunners of the more sophisticated modeling that would follow. After computing transitory earnings for the white male heads in their sample (as deviations from nine-year average earnings), they computed the cross-sectional variance of the individual transitory components in each year to look at how transitory variance changed from year to year. They found that while following a cyclical pattern, the transitory variance had trended upwards, doubling between 1973 and 1987.

In their next paper, Gottschalk and Moffitt adopted two new strategies. In what I will call their variance decomposition strategy, they modeled log earnings in each year as the sum of a permanent component (an individual fixed effect) and a transitory

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51 Moffitt and Gottschalk (1995). For methodological details see footnote 6 and the discussion in Chapter Two. Moffitt and Gottschalk (1998) later updated this paper with additional years of data.
component essentially assumed to be a random shock occurring each year without persisting into future years. Under this very simple model, the details of which I discuss in Chapter Two, the transitory variance in any year may be estimated by subtracting from the overall earnings variance the covariance between overall earnings and earnings from some other year.

Gottschalk and Moffitt showed how trends in transitory earnings variances changed using spans of different years in estimating the covariance term. In this model, the size of the span should not matter because the covariance term should always estimate the variance of an unchanging permanent earnings distribution.\(^{52}\) One can generalize the approach in a regression framework by using as individual observations covariances between current earnings and earnings in every other year.\(^{53}\) Estimating regression models such as this one, Gottschalk and Moffitt found that permanent and transitory earnings among white male heads had grown by comparable amounts over time, that both were countercyclical, and that both grew more between 1981 and 1987 than between 1969 and 1980.

Gottschalk and Moffitt modified their variance decomposition strategy slightly in their later work.\(^{54}\) Rather than specifying transitory income to be serially uncorrelated, they allowed shocks to cause deviations from permanent earnings that temporarily

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\(^{52}\) Assuming, that is, that one always looks at the same cohort of individuals, with no entry into or exit from the data over time.

\(^{53}\) The covariances are regressed on a dummy variable indicating whether or not the element is a variance, a trend variable, and an interaction between the two. The trend coefficient indicates how covariances (permanent variances) are changing over time, while the interaction indicates how the transitory variances evolve.

\(^{54}\) Moffitt and Gottschalk (2002), Gottschalk and Moffitt (2006), Gottschalk and Moffitt (2007), Moffitt and Gottschalk (2008). All of these analyses use the PSID. They all restrict the sample to prime-age male heads not in school with positive wages and either positive hours worked (2002) or positive weeks worked. All four analyses trim wages at the bottom and top of the distribution within cells defined by demographic characteristics, and all three residualize logged wages to remove the effects of demographic variables.
persist. But the revised model stipulated that if the span used in computing the covariance term was large enough, then the transitory components of earnings in the two years will be uncorrelated, so that the transitory variance may still be computed as the total earnings variance minus the covariance.

For reasons I describe in Chapter Two, Gottschalk and Moffitt no longer advocate this model, so in some sense the validity of all of their estimates that use the model might be discounted. Nevertheless, estimates using this model have figured prominently in the academic and political debates around volatility trends, so it is worth summarizing their most recent results.

Gottschalk and Moffitt’s 2006 working paper indicates that volatility among male heads was flat during the second half of the 1970s, increased during the early 1980s, and declined during the rest of the decade, returning to its 1980 level (though an up-tick mid-decade interrupted the decline). Volatility then increased again in the first half of the 1990s, fell below its 1990 level, and rose again during the latter 1990s and the early 2000s. Volatility was about 150 percent higher in 2002 than in 1974, or nearly 60 percent higher if expressed as standard deviations. A 2007 presentation by Gottschalk and Moffitt found trends generally consistent with the 2006 paper, except that after the early 1990s increase in volatility, there was little change through 1996 and a steady

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55 Gottschalk and Moffitt (2006). In this paper the authors include the SEO sample and weight the data. They restrict the sample to male heads age 20-59 not in school with positive wages and positive weeks worked. The analyses trim the top and bottom 1% of wages within age-year cells, and regress logged wages (adjusted for inflation using the CPI-U-RS) on five education categories. Gottschalk and Moffitt use four-year spans to compute covariances.
increase from 1996 to 2002. The increase in volatility is about the same as in the 2006 paper from the early 1970s to the early 2000s.

The 2007 presentation included the only results using this model with a dataset other than the PSID. Specifically, they presented a volatility trend from the Social Security Administration’s Continuous Work History Sample, estimated with the assistance of Jonathan Schwabish. These results indicated a decline in volatility from 1980 to 1989, and increase between 1989 and 1993. After a smaller decline from 1993 to 1995, volatility was flat through 1999. It then increased through 2002 before flattening out in 2003. Over the entire period, volatility was the same in 2003 as in 1980, contradicting the sizeable increases over this period in their PSID results.

The most recent results from Gottschalk and Moffitt are from a late 2008 working paper. This time using a ten-year span to compute earnings covariances, and showing results for male heads age 30-39, they find an increase in volatility over the early 1980s, followed by a decline over the rest of the decade (with a large spike in 1986). Volatility then rose in the early 1990s before falling through the late 1990s. It rose again between 1998 and 2000 but then fell. The increase in volatility from 1980 to 2004 was less than 15 percent if expressed in terms of standard deviations—much lower than in their earlier PSID results for male heads age 20-59.


57 Ibid. They residualize log earnings on age and year and trim the top and bottom 1% of observations.

58 Moffitt and Gottschalk (2008). The authors group sample members into three age groups, and compute covariance matrices within each group. Moffitt and Gottschalk use a ten-year span in computing earnings covariances. For additional methodological details, see footnote 34.
In a recent working paper, Gottschalk and Moffitt describe yet another variant on their variance decomposition strategy.\textsuperscript{59} If permanent income is conceived of in terms of human capital levels and time-varying returns to human capital, then the covariance of incomes between two years is the product of the returns to human capital in the two years and the covariance of the permanent incomes in the two years. Taking the natural log of both sides provides an equation relating the log covariance of income to the sum of the logs of the returns to human capital and the log of the covariance of permanent income. One can estimate this equation with OLS using year dummies and approximating the last term using some function of age and lag length between incomes.\textsuperscript{60} Using the estimated returns to human capital and the estimated covariance of permanent income between ages $a$ and $a-0$ (i.e., the estimated permanent variance at age $a$), the product of the return to human capital at time $t$, the return to human capital at time $t-0$, and the covariance of permanent incomes measured at ages $a$ and $a-0$ equals the permanent variance at age $a$ in year $t$. Subtracting this from the total earnings variance for adults of age $a$ in year $t$ gives the transitory variance of adults of age $a$ in year $t$.

The trend in these transitory variances is rather different from the trend produced by simply subtracting the covariance of earnings 10 years apart from the total earnings variance. Volatility declined over the early 1970s, then increased unevenly through the mid-1980s. It showed no consistent trend during the second half of the 1980s but increased again in the early 1990s. Volatility then declined through 1998, increased

\textsuperscript{59} Ibid.

\textsuperscript{60} The log of covariances at different lags (of length 10 and greater) is regressed on year dummies (which equal 1 if the year is included in a covariance) and second-order polynomials in age and lag length.

The third strategy Gottschalk and Moffitt have adopted in their research is to estimate more complex error components models.61 These models specify different structures relating the permanent and transitory components of earnings to their past values. In the model on which Gottschalk and Moffitt have settled, today’s permanent earnings are specified to depend on the returns to human capital, which is itself dependent on yesterday’s innovation to human capital, an individual’s rate of steady human capital growth, and a shock to human capital. Today’s transitory shocks persist into tomorrow, but weakening over time (at different rates from year to year), with the variance of the shocks varying over time. Technically, the model specifies that permanent earnings evolve by a random walk with a random growth factor and that transitory earnings—the variance of which changes over calendar time—evolve according to an ARMA(1,1) process.

Their most recent results indicate that among 30- to 39-year-old male heads, volatility declined in the early 1970s, then rose unevenly through the mid-1980s.62 There was little consistent trend over the rest of the 1980s, but volatility increased again in the early 1990s before declining through the late 1990s. It increased again between 1998 and 2002 and then declined from 2002 to 2004. From 1973 to 2004, volatility increased by about two-thirds when expressed in standard deviations, loosely following a countercyclical pattern.


The only subsequent research to use Gottschalk’s and Moffitt’s first strategy estimating transitory variances was Kopczuk et al. (2009). They measured volatility by computing deviations from a five-year average of earnings for each person, and then looking at the cross-sectional variance of deviations across people. They found that volatility rose in the early 1940s, fell through the early 1960s, and changed only modestly thereafter.

Similarly, only one other study has utilized Gottschalk’s and Moffitt’s second strategy. Jacob Hacker and Elisabeth Jacobs estimated covariance terms in each year using male heads’ labor income measured four years apart and then subtracted the covariance from the earnings variance to estimate transitory variances. They found an increase in volatility during the recession of the early 1970s, flat volatility the rest of the decade, and sharply increasing volatility during the early-1980s recessions. Volatility then declined through the mid-1980s before flattening out. The 1990s repeated this cyclical pattern, with a large increase in volatility followed by an almost equally large decline, followed by another sharp increase after 2000 and a smaller decline. The increase from 1974 to 2002 was even higher than Gottschalk and Moffitt’s estimate: 210 percent, though that drops to around 75 percent if the results are expressed in standard deviations.

Most subsequent volatility research looking at dispersion of earnings shocks has relied on different versions of Gottschalk’s and Moffitt’s error components strategy. Nearly as oft-cited as Moffitt and Gottschalk’s results are those of Steven Haider. Haider

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63 Hacker and Jacobs (2008). The authors use the SEO sample and weights, and they restrict the sample to male heads between the ages of 25 and 62.
examined the transitory variance of white male wage and salary income from 1967 to 1991 using the PSID. He found a doubling of volatility over that period. Like Moffitt and Gottschalk, Haider found an increase in volatility during the 1970s, though the trend along the way looks quite different. Haider found little change in volatility during the 1980s rather than an increase. Both studies found a sharp increase in volatility from 1990 to 1991. Overall, Haider’s estimates imply a bigger increase in volatility than Moffitt and Gottschalk’s do, largely due to a bigger increase in the 1970s. His results also show a strong countercyclical pattern.

A recent study by Mary Daly and Robert Valletta used Haider’s model to estimate volatility trends in the U.S., Great Britain, and Germany. Daly and Valletta found the same basic trend as Haider among American white male heads from 1979 to 1982, but while Haider found a slight decline over the 1980s, they found a slight increase. The increase accelerated in the early 1990s before volatility declined between 1992 and 1996. Overall, the increase from 1979 to 1996 was about 25 to 30 percent.

Ann Huff Stevens estimated trends in male heads’ transitory wage variances using an error components model and the PSID. She found that volatility declined in the

64 Haider (2001). After regressing log wages on a quartic in experience, Haider fits his model, which includes an individual fixed effect, individual-specific slopes with respect to experience, a time-varying factor loading that applies to both, and a transitory component that evolves according to an ARMA(1,1) process. For additional methodological details see note 14.

65 Daly and Valletta (2008). The authors use the Cross-National Equivalent File, a multi-country dataset that includes variables for the U.S. created from the PSID (see Burkhauser and Lillard, 2007). Daly and Valletta examine white male heads between the ages of 25 and 61. They exclude nonpositive earnings, as well as men with $0 reports between two years with positive earnings, and they trim the top and bottom 1 percent of positive earnings.

66 Stevens (2001). Stevens first regresses log earnings on race, education, and potential experience. She uses the covariance matrix of the residuals to estimate the parameters in her model, which includes a person-specific fixed effect with a time-varying factor loading, and a transitory component that evolves according to an AR(1) process. She includes male heads age 24 to 64 with nonzero wage and salary earnings.
early 1970s, increased from 1974 to 1976 before declining again, and rose from 1978 to 1983. Volatility then declined through 1987 before increasing through 1991. The trend differs from Moffitt and Gottschalk’s in that Stevens’s estimates are relatively flat from 1972 to 1987 rather than increasing steadily. Again, however, a countercyclical pattern is evident.

While all of these error components models rely on the PSID, Bhashkar Mazumder estimated trends in transitory variances using several SIPP panels matched to Social Security Administration earnings data.67 His model indicates that earnings volatility among men declined from 1984 to 1989 (with a spike in 1988), increased through 1995, and then was flat through 1997. Expressed in terms of standard deviations, the increase from 1984 to 1997 was about 25 percent. For comparison, Gottschalk’s and Moffitt’s estimates, as well as Daly’s and Valletta’s, show very little change between 1984 and 1997.

Rather than modeling trends in the volatility of annual earnings, as all of these studies do, Dean Hyslop looked at trends in the transitory variance of hourly wages.68 Hyslop’s analyses are confined to continuously married couples, but they include separate trends for men and women. The transitory variance of wages increased

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67 Mazumder (2001). Mazumder de-meaned earnings in each year within each birth cohort. He then modeled de-meaned earnings as a function of a permanent component (featuring an individual fixed effect, an individual deviation from the birth cohort’s growth rate, and a random walk) with a time-varying factor loading and a transitory component (following a first-order autoregressive process with innovations modeled as following a quartic in experience and with a time-varying factor loading on innovations). For additional methodological details, see note 21.

68 Hyslop (2001). Hyslop also used the PSID, including the SEO sample. His sample includes continuously married couples age 18-60 in 1980 with no top-coded data and no wages higher than $100 an hour. Both must have positive earnings and hours in every year. Wages (labor income divided by hours worked) are adjusted for inflation using the CPI. Husbands’ and wives’ wages are modeled as functions of average wages for the population, individual-specific permanent components with time-varying factor loadings, time-varying transitory components subject to factor loadings, and measurement error. Transitory components are modeled as following a first-order autoregressive process, and husbands’ and wives’ permanent and transitory components are allowed to be correlated.
(unevenly) from 1979 to 1985 among both, rising just 15 percent among husbands but doubling among wives.

Marco Leonardi also estimated an error components model based on hourly wages, using the PSID and focusing on all male heads. He found that the transitory variance of wages increased about 60 percent from 1969 to 1991, according to the line he fits to his estimates, which shows the transitory variance climbing at an accelerating rate from the early 1980s forward.

Finally, there are four studies that model earnings as subject to both transitory and permanent shocks. Costas Meghir and Luigi Pistaferri estimated a model using the PSID that attempts to distinguish the transitory component of earnings from measurement error. Using estimates from previous studies, they assumed that measurement error accounts for 25 percent of the variance of the rate of earnings growth. Pooling all years, their model implies that measurement error overstates transitory variances by 27 to 62 percent (with the effect increasing with education). They found that the variance of transitory shocks was relatively flat from 1967 to 1970, rose from 1970 to 1972, and fell again in 1973 before increasing to a new high by 1975. Volatility then fell back to earlier levels by 1977. There was no trend from 1977 to 1981, but volatility jumped in 1982. It

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69 Leonardi (2003). In his model, permanent earnings evolve by a random walk, with a time-varying factor loading, and transitory earnings evolve according to an AR(1) process, also with a time-varying factor loading. He first regresses earnings on a quadratic in experience. Leonardi excludes the SEO sample of the PSID, examines male heads 20-59 years old, and excludes students and the self-employed. He also excludes those with topcoded earnings, those with a wage less than half the minimum wage, and those working fewer than 520 hours or more than 5,096 hours.

70 Meghir and Pistaferri (2004). The authors use the PSID to look at male heads’ labor income volatility, focusing on those age 25 to 55 who have at least nine years of labor income data. They exclude individuals with missing values on covariates and who experience very large earnings changes. As a first step, they regress log earnings on year dummies, a quadratic in age, and race, region, and urbanicity dummies. These regressions are done separately for each of three education groups. They then use the residuals to fit their model, which specifies that permanent income follows a random walk, that transitory income evolves by a MA(1) process, and that classical measurement error affects results in all years.
then declined through 1987 (to a new low). Volatility remained low through 1992. Over the entire period, volatility increased by 40 percent (expressed in standard deviations), though the increase is dwarfed by the swings over time.

Meghir and Pistaferri estimated that the variance of *permanent* shocks to earnings declined in the early 1970s, increased notably through the early 1980s, declined sharply through mid-decade, and increased somewhat over the late 1980s and early 1990s. From 1969 to 1991, the increase was 25 to 30 percent.

Shane Jensen and Stephen Shore also use the PSID to estimate a model with permanent and transitory shocks.71 Their approach is unique in that it attempts to model variances of individuals’ distributions of permanent and transitory shocks from which their actual shocks are drawn. They then report the mean and median of the individual variances. Their results imply that the mean variance of male heads’ transitory earnings shocks rose in the early 1970s, then fell, rose, and fell again over the rest of the decade. It rose in the early 1980s and showed little trend through 1991. After increasing in 1992, the transitory variance declined over the rest of the decade, rose between 2000 and 2002 and then fell again. Expressed in standard deviations, the increase from 1969 to 2004 was about two-thirds. On the other hand the median transitory variance was flat over the period, implying that the increase over time was driven by rising volatility among those with relatively high volatility.

The median variance of *permanent* shocks was also flat even as the mean variance increased. The trend in the mean variance of permanent shocks was roughly similar to

71 Jensen and Shore (2008). Income is modeled as a function of (1) demographics, (2) a permanent component that is the sum of initial income plus past permanent shocks weighted such that their impact takes time to fully set in, and (3) a temporary component that is the sum of past temporary shocks weighted such that their impact gradually disappears. For additional methodological details, see note 26.
that for transitory shocks, except that the increase over the 1980s was smaller and volatility of permanent shocks grew in the 1990s while transitory volatility shrank slightly. Over the entire period, the variance of permanent shocks increased by about 50 percent.

Fatih Guvenen’s recent paper is unique in presenting trend estimates based on two distinct models, one of which has random effects in income trajectories (the heterogeneous income profiles, or “HIP”, model) and one of which includes only individual fixed effects (the restricted income profiles, or “RIP”, model). Guvenen presented evidence that the HIP model is more appropriate. He found that according to the HIP model, among male heads the transitory variance shows little trend from 1968 through 1985 and then increases erratically between 1985 and 1992. The increase over the entire period was about 23 percent when volatility is expressed as standard deviations. According to the RIP model, the transitory variance increased unevenly from 1968 to 1992, with volatility rising 28 percent when expressed as standard deviations.

Permanent volatility, based on the HIP model, showed little trend through the late 1970s but rose through the early 1980s. It then declined through 1990 before increasing again, rising 15 percent between 1968 and 1992. The RIP model showed permanent volatility following little consistent trend through the mid-1970s, increasing through the mid-1980s, then declining through the early 1990s. Overall, the increase was 33 percent from 1968 to 1992.

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72 Guvenen (2009). He uses the PSID, excluding the SEO sample. The sample includes male heads age 20 to 64 who (for at least 20 years) have positive labor income and hours, annual work hours between 520 and 5110, and an hourly wage equivalent to $2 to $400 in 1993. Labor income is modeled as a function of a cubic in potential experience that is common to all heads but time-varying, individual fixed effects, individual-specific linear income profiles, a component following an AR(1) process with an innovation that has a time-varying variance, and a transitory shock that also has a time-varying variance.
The last paper with an error components model that includes permanent shocks examines volatility of hourly wages. Jonathan Heathcoate and his colleagues estimate trends in the variance of shocks among men and women in continuously married households. Their estimates fluctuate annually but show an increase over the 1970s and 1990s, with little change during the 1980s. From 1967 to 2000, the transitory variance increased by about 50 percent, expressed in standard deviations. The variance of permanent earnings shocks also increased in the 1970s and 1990s, declined in the 1980s, and increased about two-thirds over the entire period.

While there are many inconsistencies across the studies in this line of research, some generalizations are possible. Volatility was either flat or increasing over the 1970s. Most studies find rising volatility in the 1980s, though several find a flat or declining trend. Most notably, two studies that use Social Security Administration data rather than the PSID show a decline over all or part of the 1980s (Gottschalk and Moffitt, 2007; Mazumder, 2001). The research generally agrees that volatility rose in the early 1980s recession and declined in the mid-1980s, though the transitory volatility estimates of Guvenen show the opposite pattern. The research also consistently shows that volatility increased during the recession of the early 1990s and over the course of the decade. The evidence is inconsistent as to whether it increased in the early 2000s. Among the studies that span the years in the PSID, male earnings volatility increased by 33 to 165 percent.

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73 Heathcoate et al. (2008). The authors use the PSID, excluding the SEO. They measure hourly wages as annual labor income divided by annual hours and adjust earnings using the CPI. The sample is confined to adults in married households where the husband is 25-59 and worked at least 260 hours, where both spouses earned more than half the minimum wage (if they worked), and where there is no self-employment income. Log hourly wages are modeled as a function of a permanent component that follows an autoregressive process, a transitory component, and measurement error.
from the early 1970s to the early 2000s (15 to 65 percent expressed as standard deviations. Kopczuk et al.’s results imply that volatility was higher before 1960 than it was in subsequent decades.

Other Approaches to Volatility Measurement

The papers by Buchinsky and Hunt, Kopczuk et al. (2009) and by Fields and his colleagues present estimates of short-term earnings mobility trends that have no equivalent in other research. Their measure is based on a comparison of inequality when a multi-year average of earnings is used versus the average inequality across each individual year. When there is substantial mobility from year to year, inequality will be smaller using a multi-year average than in any individual year. The measure formalizes this tendency in a single number. Kopczuk et al. report that earnings mobility for commerce and industry workers, and for men in particular, rose from 1940 to 1945, fell over the next few years, and then gradually declined until the early 1960s. It rose a bit over the next decade, but beginning in the early 1970s mobility gradually fell again, ending lower in the early 2000s than in any previous year.74

Fields and his colleagues found that male labor income mobility increased in the 1970s, was flat from the late 1970s to the early 1980s, and declined over the 1980s and early 1990s.75 Buchinsky and Hunt found that mobility in hourly wages and in annual wage income declined steadily over the 1980s, except that when those with no wages

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74 Kopczuk, Saez, and Song (2009). The authors compare the Gini coefficient using a five-year earnings average to the average of the five individual Gini’s. For additional details, see note 4.

75 Fields, Leary, and Ok (2000). Their measure is 1 minus the ratio of an inequality measure using the average of earnings over a window of six years to the average of the same inequality measure across the individual years within the window. The inequality measure is the Gini coefficient. For additional methodological detail, see footnote 7.
were included, mobility in hourly wages was flat or slightly increasing in the latter part of the decade.\textsuperscript{76} The authors report that the results were similar for men and women.

Fields and his colleagues also used a second unconventional measure.\textsuperscript{77} They computed for each male in their sample the absolute value of the change (over five years) in the individual’s share of aggregate income and then reported the mean of this number across men. Short-term mobility defined in this way increased over the 1970s and through the early 1980s, was flat over the 1980s, and declined from the late 1980s to the early 1990s.

As part of their examination of mobility and volatility trends, Daly and Duncan used a measure that defined volatility as the median across workers of the within-person average over 11 years of year-to-year unsigned percent changes.\textsuperscript{78} Consistent with their other measures, they found that labor income volatility among men was higher in the 1980s than the 1970s.

\textbf{Summary of Previous Research on Earnings Mobility and Volatility}

Summarizing the literature on directional short-term mobility, it appears that there were large declines in upward and downward relative mobility from the postwar years through the mid-1960s. Similarly, downward absolute mobility declined through the first part of the decade. In the late 1960s, relative mobility probably increased to its 1960

\textsuperscript{76} Buchinsky and Hunt (1999). The measure of mobility is 1 minus the ratio of an inequality measure using the average of earnings over windows of different duration to the average of the same inequality measure across the individual years within the window. The inequality measure is a generalized entropy index that considers each person’s share of total earnings against a perfectly equal distribution. For additional methodological details, see footnote 10.

\textsuperscript{77} Ibid.

\textsuperscript{78} Daly and Duncan (1997). For methodological details, see footnote 8.
level, and downward absolute mobility also increased (still ending the decade lower than when it started).

From the 1970s onward, a countercyclical pattern in directional (relative or absolute) mobility prevailed, with downward mobility temporarily increasing during recessions and upward mobility temporarily decreasing. During the 1970s directional mobility changed little among all adults, though this aggregate trend masks an increase in downward absolute mobility among men and a small decline among women.

In the 1980s, directional relative mobility declined, and absolute mobility either was unchanged or declined. In the 1990s, relative mobility changed little, and if anything, fewer people experienced downward absolute mobility while more had upward absolute mobility. The first years of the current decade saw little change in relative mobility but deterioration in absolute mobility, with increases in downward mobility and declines in upward mobility, thanks to the 2001 recession and the weak recovery. The research on directional mobility is fairly consistent then, and indicates no detrimental secular trend in earnings movements since the 1970s, and only among men in that decade.

Research on nondirectional relative and absolute mobility, on intertemporal earnings association, and on dispersion in earnings changes might be expected to show similar trends, since all of these measures summarize one-way movement between two points in time without regard to the direction of movements. Instead, the main consistency across these literatures is inconsistency within each research area. Only the research on nondirectional relative mobility shows internally consistent results over the
last 40 years as a whole, indicating a probable increase in earnings movement over the 1970s followed by flat or declining trends in subsequent decades.

Only for the 1970s does the research across these literatures allow for a reliable conclusion about a decade, implying that the trend in earnings movement was increasing or flat for men. In the 1980s, there is fairly consistent evidence that earnings movements declined at least later in the decade, and the weight of the evidence in all but the dispersion-of-changes literature indicates a decline over the decade (and that literature indicates that earnings movement declined for women over the decade). Research on male earnings movements for the 1990s and early 2000s in these four areas is either internally inconsistent (absolute mobility, intertemporal association for the 1990s), conflicting (relative mobility and dispersion of changes), or nonexistent (intertemporal association for the 2000s). The absolute mobility and dispersion-of-changes literature implies declines in earnings movement among women in the 1990s.

Perhaps a reasonable conclusion from this body of work is that at least since the 1980s, secular trends in nondirectional earnings movements among men have been small enough that different researchers—often using the same dataset and sometimes using very similar measures—have reached different conclusions. That would be the case if, for example, trends in upward and downward mobility have largely canceled each other out.

That said, if the Social Security Administration data is more reliable than the PSID, then the evidence implies that earnings movements declined in the 1980s and were then constant or declining through the 1990s among men, declining among women. In the early 2000s, it increased or was flat among men and was flat or declining among
women. Whether one prefers the SSA data or the PSID, however, there is little evidence of detrimental trends in earnings movements after the 1970s.

The research on within-person earnings dispersion and on the dispersion of earnings shocks more clearly measures “volatility”. Volatility as measured by within-person dispersion increased in the 1970s and probably increased among men through the 1980s or early 1990s. The trends across the 1980s and the 1990s are generally inconsistent however. Volatility probably declined in the early 2000s but likely remained higher than in 1980 for men. There is some evidence that among women volatility declined after 1980, though the research is sparse.

Volatility as measured by the variance of shocks to earnings increased or was flat during the 1970s. Most studies show an increase in the 1980s, though the research that uses the Social Security Administration data implies declines. Volatility increased further in the 1990s; the evidence for the early 2000s is inconsistent. As with the within-person-dispersion research, volatility levels were generally higher among men in the early 2000s than in 1980, though the research differs as to the timing of increases.

In sum, the 1970s appear to have been a bad decade for men in terms of downward earnings mobility and volatility, and volatility may have also increased among women. Since 1980, downward and upward mobility have followed a cyclical pattern, and the secular change in directional mobility has been minimal. However, volatility as measured by within-person dispersion of earnings or by dispersion of shocks to earnings appears to have increased between 1980 and the early 2000s.

This conclusion, however, relies on evidence that is or may be incomplete in several ways. First, the PSID-based studies—which constitute much of the research,
which rely on a survey with a number of features that greatly complicate its use in research on volatility and mobility, and which often has produced findings that differ from those obtained from Social Security Administration data—may not accurately portray trends in earnings instability. This could be true either because of problems with the PSID data itself or with the methodological decisions past researchers have made. My analyses in Chapter Two attempt to improve the methodological decisions used by previous researchers and to standardize them across different types of mobility and volatility measures.

The second question is whether the within-person dispersion measures typically used or the model-based measures of dispersion of shocks adequately differentiate between trends in earnings growth and trends in volatility per se. Finally, the existing research is sparse in certain aspects. Most obviously, there is much less research on trends among women than among men. Furthermore, for some measures there exist no time series that include estimates for men and women separately as well as combined. In addition, some of the literatures do not extend as far back as the data allows or as far forward. The estimates in Chapter Two are an attempt to fill in these gaps.
Appendix Two:

Previous Research on Income Instability and Volatility Trends

Unlike the research on earnings volatility and instability, most of the studies on family and household income volatility have been produced in the past five years. To a remarkable extent, the development of this literature has taken place not in academic forums, but within think tanks, government agencies, mass-marketed books, and even newspaper columns and outlets of opinion journalism.

Unfortunately, because the earliest figures to garner wide attention turned out to be problematic, it has been unclear what conclusions should be drawn from the research on income volatility trends. In this appendix, I summarize the findings across a range of studies of income instability and volatility. To organize the literature, I return to the categorization from Appendix One’s literature review. Chapter Three presents my own results from the PSID that correspond to each of these categories.

Research on Short-Term Relative Mobility

I begin with research on short-term relative mobility, which looks at changes in ranks over periods of five years or less. Unlike the research on relative earnings

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1 I will generally use “family income” and “household income” interchangeably in this review. The PSID asks about the incomes of “family unit” members. Family units in the PSID include most cohabiting partners as well as relatives of the family unit head and his cohabiter (“heads” in the PSID are almost always men). There are only a small number of households in the PSID with multiple family units interviewed.

2 As in Appendix One, I restrict my review to studies of volatility and instability in the United States. For international evidence, see Morissette and Ostrovsky (2005) on Canada; Burkhauser et al. (1997) and Maasoumi and Trede (2001) on Germany; and Blundell and Preston (1998) on Great Britain.
mobility, previously published time series on relative income mobility extend only as far back as the late 1960s. Peter Gottschalk and Sheldon Danziger examined relative mobility in the PSID among prime-age individuals, adjusting income for family size.\(^3\) Non-directional mobility, as measured by 100 minus the percent staying in the same quintile in adjacent years, showed little trend between 1969 and 1976. It declined from 1976 to 1980. Over the 1969-1980 period, mobility changed very little.

Gottschalk and Danziger found that upward mobility (100 minus the percentage remaining in the bottom quintile) increased between 1969 and 1978 and declined from 1978 to 1980, increasing over the entire period. Downward mobility from the top quintile showed little trend from the late 1960s through the mid-1970s but declined from 1976 to 1980. Downward mobility declined over the 1970s as a whole.

Maury Gittleman and Mary Joyce also examined trends in the probability of remaining in the same family income quintile over successive years using the PSID.\(^4\) They found that non-directional mobility increased from 1968 to 1976 and then fell a bit late in the decade. Looking at the same measure but comparing incomes separated by five years, they found an increase in mobility from 1972 to 1979, with the decade again finishing with declining mobility.

\(^3\) Gottschalk and Danziger (1998). The authors confine their analyses to individuals age 22-62 in both years within a pair. They adjust incomes for family size by dividing by the poverty line. The SEO sample is excluded from these analyses.

\(^4\) Gittleman and Joyce (1999). The authors include in their sample heads and partners between the ages of 25 and 64, using the SEO and weighting. Their measure is family income adjusted for family size by dividing by the poverty line. They exclude those with incomes below $1,000 or above the lowest real topcode across years.
Joel Slemrod used administrative data from the IRS to estimate downward mobility from top deciles. He found that downward mobility from the top 5% or top 10% of pre-tax income increased from 1968 to 1969 and then declined through 1972. Downward mobility from the top 1% increased from 1968 to 1971. In 1973, downward mobility from the top 5% or top 10% increased, finishing very close to its 1968 level. Downward mobility from the top 1% declined from 1971 to 1973. Finally, downward mobility from the top 5% and top 1% increased between 1973 and 1980, exceeding all earlier levels in the latter year. Downward mobility from the top 10% probably increased a bit as well.

Peter Gottschalk, Sara McLanahan, and Gary Sandefur used the PSID to examine trends in the likelihood of staying in the bottom family income quartile in two adjacent years among all people in the PSID. If one measures upward mobility as one minus this proportion, the trend was basically flat during the 1970s, staying at or just over 20 percent.

In the 1980s, Gottschalk and Danziger found non-directional mobility continuing to decline through 1991. Upward mobility continued to decline between 1980 and 1988, then increased through 1991. It was a bit higher in 1991 (25%) than in 1969 (22%).

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5 Slemrod (1992). Slemrod technically measures the share of adults in top deciles that were not there in the previous year, but since he uses balanced panels, this is equivalent to the share of adults in the top decile the previous year who are no longer there in the current year. He relies on two different panels of tax filers – one from 1967 to 1973 and the second from 1979 to 1986. He uses all tax filers in his sample (tax returns being the unit of analysis). That means his analyses exclude those who have no income tax liability and those who do not file. The panels are based on the Social Security number of the primary taxpayer, so there is attrition when that number changes due to marriages or a couple switching who the primary taxpayer is. Very late filers are also attritors. Previous research has shown attritors to be poorer than other sample members, to be less likely to be married, and to be younger. Slemrod’s income measure is adjusted gross income plus adjustments plus excluded dividends and long-term capital gains.

6 Gottschalk, McLanahan and Sandefur (1994). The authors do not elaborate on the methodological details of their study.
to 1989. It declined from 1989 to 1991, finishing lower (21%) than its 1969 level (26%).

Across Gottschalk and Danziger’s measures, mobility consistently declined from the mid-1970s to the early 1990s, though not smoothly. Gottschalk, McLanahan, and Sandefur found that upward mobility continued to be flat throughout the first half of the 1980s.

Gittleman and Joyce found a decline in non-directional mobility from 1979 to 1991, with mobility a bit higher in 1991 than in the late 1960s. Their five-year mobility measure indicated a decline between 1979 and 1985. Mobility then increased sizably from 1985 to 1987, followed by a large decline between 1987 and 1988. After a small increase in 1989, mobility was flat through 1991, ending at a level similar to the early 1970s and early 1980s levels.

Slemrod found little trend in downward mobility from the top 5% and top 10% from 1980 to 1984, but it increased from 1984 to 1986. Downward mobility was higher in 1986 than in 1980, increasing from 22 percent to 27 percent for the top decile and from 26 percent to 31 percent for the top 5 percent. On the other hand, downward mobility from the top percentile increased from 28 percent in 1980 to 33 percent in 1982, was flat for several years, and then increased to 40% in 1986. Slemrod noted that the increases in downward mobility at the end of the period were at least partly due to taxpayers realizing capital gains in anticipation of tax law changes taking effect in 1987.

Also using administrative records from tax returns, Robert Carroll and his colleagues examined non-directional mobility in the 1980s and early 1990s. They

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7 Carroll et al. (2007). The authors’ sample included primary taxpayers at least 30 years old in 1979 and under 62 years old in 1995. Taxpayers who did not have returns in all 17 years were excluded. Their income measure – “constant law income” – is fairly comprehensive. The authors also examined mobility over 3- and 4-year periods, which show trends broadly consistent with the 2-year trends.
showed that in 1980, 34 percent of primary taxpayers age 30-46 the previous year had changed income quintiles. This share declined to 29 percent by 1983, showed little trend through 1988, and then declined further the rest of the decade. The share of primary taxpayers rising up from the bottom quintile was flat over the 1980s, and the share falling down from the top quintile declined a bit.

Greg Duncan, Tim Smeeding, and Willard Rodgers defined “low”, “middle”, and “high” income groups based on real income cutpoints and then held the size of those groups constant over time. They found that upward mobility from the “low” income category to the “middle” or “high” categories was lower during the 1980-87 period than during the 1967-1980 period. Similarly, downward mobility from the “high” category was also lower. Downward mobility from the “middle” category increased, however, and upward mobility from the “middle” was unchanged.

In the 1990s, Carroll et al. found an increase in non-directional mobility from 1990 to 1995. Upward mobility from the bottom quintile and downward mobility from the top were both flat.

Summarizing these studies on short-term relative income mobility, relative income mobility declined or was stable between the late 1960s and the mid-1990s, providing little evidence of a Great Risk Shift. Non-directional mobility either declined or was stable in the 1970s and declined in the 1980s. Carroll and his colleagues found an increase in mobility over the first half of the 1990s, though mobility remained below its

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8 Duncan, Smeeding, and Rodgers (1993). The authors used the 1968-1987 waves of the PSID, applying weights and using the SEO sample. The men and women in the sample were initially aged 25-50. Household income was defined to include food stamps and to exclude federal taxes, and incomes were adjusted for inflation using the CPI-U-X1. The authors examined transitions looking at income averaged over years \( t \) and \( t+1 \) and income averaged over years \( t+3 \) and \( t+4 \). The three income groups created constituted 18%, 73%, and 9% of households (low, middle, and high, respectively).
Gottschalk and Danziger found declines in downward mobility from the top quintile in the 1970s and 1980s, and Carroll et al. confirmed the 1980s decline. Slemrod, however, generally found increasing downward mobility from the top decile, ventile, and centile. Carroll et al. found no change in downward mobility in the early 1990s. Finally, upward relative mobility from the bottom quintile either was unchanged or declined in the 1970s, in the 1980s, and in the first half of the 1980s in particular. Carroll et al. again found no change in the early 1990s.

Relative income mobility trends are similar to relative earnings mobility trends in some aspects but different in others. While it appears that non-directional earnings mobility increased in the 1970s, family income mobility was either stable or declined. Among all adults, male and female, upward and downward earnings and income mobility changed very little, though directional income mobility of both types may have declined, as did upward earnings mobility in the latter part of the decade. Relative mobility, directional or not, in terms of either earnings or income, appears to have declined in the 1980s. The first half of the 1990s saw declines in non-directional earnings mobility but increases in income mobility. Upward and downward mobility, either in terms of earnings or income, however, appear not to have changed much. Finally, trends in directional income mobility do not display the same countercyclical pattern exhibited by the directional earnings mobility trends. This important result would seem to imply that to a large extent, families compensate for the earnings changes of individual members by adjusting the work decisions of other family members or through public and private transfers.
Research on Intertemporal Income Associations

In contrast to research that includes correlation matrices for individual earnings over time, because there has been less research modeling household income dynamics, intertemporal income correlation matrices are far less common. I found just one study that examined trends in short-term intertemporal income associations. Tom Hertz examined household income mobility using matched CPS files. He found mobility as measured by one minus the correlation over adjacent years increased from 1991 to 1998. It then declined slightly in 2004, though the change was not statistically significant. The 1990s trend is the opposite of what Bhashkar Mazumder (2001) found for men’s earnings, but these are the only two comparable studies between the earnings and income literatures using correlations.

Research on Short-Term Absolute Mobility

Rather than defining mobility in terms of movement in ranks or in terms of intertemporal association, it can be conceptualized as movement in absolute income levels. Beginning in 2006, Hacker’s research began to incorporate short-term absolute mobility measures. As with the transitory variance measures, Hacker revised his original estimates downward after GRS’s publication, which claimed that the typical two-year income drop among families experiencing a drop amounted to 44 percent of income in 1998 (up from 27 percent in 1971). The revised edition of GRS did not mention the magnitude of the typical drop, so these figures should presumably be disregarded.

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9 Hertz (2006). The data is from the March supplement to the CPS, using three pairs of years. The matched subsamples are re-weighted to be representative of the first year in each pair. Incomes are adjusted for inflation using the CPI-U-RS.
The new *GRS* analyses indicated that for the years 1998-2000, 2000-2002, and 2002-2004, an average of 45 percent of nonelderly adults experienced a two-year drop in family income and that this percentage, while fluctuating with the business cycle, had “remained fairly steady” since the 1970s.10 Roughly 4 percent of adults experienced a drop in income of at least 50 percent between 1969 and 1971, a figure that nearly doubled by 1981, fluctuating with the business cycle along the way. After falling to about 5 percent by 1984, downward mobility was flat through 1989. It then rose sharply to 9.5 percent by 1993, fell to 6.5 percent by 1998, increased nearly to 10 percent by 2002, and finally fell back to about 8 percent in 2004.

Jacobs found patterns very similar to Hacker’s revised figures in her initial analysis, showing the same trends in the percentage of people experiencing a 50 percent drop over two years, but lower levels and a smaller increase over time.11 She omitted the 1991-95 data points due to concerns about the validity of the estimates in those years. Nevertheless, Hacker’s and Jacobs’s subsequent research brief for the Economic Policy Institute reproduced Hacker’s *GRS* chart almost exactly.

Hacker also indicated in a footnote of *GRS* that 15 percent of adults experienced a 50 percent drop over two years at least once in either 1971-73, 1972-74, or 1973-75,

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10 Hacker (2008). Some information on the methods is available at http://pantheon.yale.edu/~jhacker/method.html. Hacker used the pre-tax family income variable in the PSID and unlike in Hacker (2006), he did not rely on the CNEF at all in his revised analyses. He adjusted income for family size by dividing by the square root of family size, and he then logged the measures and adjusted them for inflation (using the CPI-U this time). His sample was confined to adults 25-61, including the entire core sample. This time he applied the sample weights. Hacker dropped all incomes of $1 or less and then trimmed an additional 1 percent of observations from the bottom, and he top-coded incomes at the top-code level for the year in which the most incomes are affected.

11 Jacobs (2007). Jacobs used the PSID, examining the pre-tax family income (adjusted for family size and inflation) of persons age 25-61. She trimmed the top and bottom 3 percent of incomes.

Finally, Hacker reported trends in the predicted probability of a drop in income in several outlets. To do so, he ran logistic regression models that included a time trend, individual fixed effects, and controls for age, education, race, gender, income averaged over five years, and dummy variables for having experienced each of several income-affecting events. He then reported the predicted probabilities of losses of a given size holding all control variables at their average values.

The predicted probability of a 20% drop in income between two years increased from 4 percent to 11 percent between 1970 and the early 2000s, with the increase concentrated in the early 1970s, early 1990s and early 2000s.13 The predicted probability of a 50% drop in income between two years increased from 7% to nearly 17%, and the increase was fairly steadily over the entire period, with an anomalous bump in the early and mid-1990s.14 Clearly these two sets of figures are inconsistent, since the predicted probability of a 20% drop should be higher than the predicted probability of a 50% drop, and Hacker omitted these analyses from later volatility publications.

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12 Hacker (2008), footnote 41 of Chapter 1.

13 Hacker (2005). See also Hacker (2007c). Hacker does not indicate what specific variables are used in his logistic regression models.

14 Hacker (2006). Hacker used the PSID and a version of the PSID called the Cross National Equivalent File that includes modified income variables and tax estimates. His measure of pre-tax family income was a composite variable that included taxable income, public energy subsidies, and the rental value of free housing from the PSID, and public and private transfers from the CNEF. The post-tax measure subtracted federal and state income taxes, and payroll taxes taken from the CNEF, and property taxes taken from the PSID. Hacker adjusted income for family size by dividing by square root of family size, and he then logged the measures and adjusted them for inflation using the International Monetary Fund CPI (from the CNEF). His sample was confined to adults 25-61, including the entire core sample. He did not, however, apply the sample weights. These details were obtained through personal communication with Hacker in 2006 and 2007. Footnote 40 in Hacker (2006) gives the events included in the model: unemployment, retirement, disability, illness, divorce, marriage, birth of child, and adoption of child.
Reporter Peter Gosselin’s research on income volatility has rivaled Hacker’s over the past year in the popular attention it has received. Gosselin’s work on the topic began in a series of long front-page articles in the *Los Angeles Times*.\(^{15}\) It culminated in the publication of *High Wire: The Precarious Financial Lives of American Families* in 2008. Gosselin reported that the share of adults in the PSID experiencing a 50% drop in family income from one year to the next grew from just 3 percent in the early 1970s to 6 percent in the mid-1990s. The share experiencing a 50% drop over two years rose steadily from 5 percent in the 1970s to 9 percent in the early 2000s.\(^{16}\)

Like Hacker, Gosselin also produced a complementary set of analyses to look at changes in the likelihood of experiencing different income-affecting events and of experiencing large income drops as a consequence of them. The risk of a 25% drop in income between two years rose from 17 percent between 1974 and 1983 to 19 percent between 1994 and 2003, while the risk of a 50% drop rose from 5 percent to 8 percent. Using figures in Gosselin and Zimmerman (2008), one can determine that the risk of experiencing any of seven income-affecting events \textit{and} a 50% income drop grew from 3 percent to 4 percent between 1974-83 and 1994-2003.\(^{17}\) Over time, these events came to

\(^{15}\) Gosselin (2004a, 2004b, 2004c).

\(^{16}\) Gosselin (2008). Gosselin used the PSID and focused on adults 25-64 years old. He adjusted family incomes for inflation using the CPI-U-RS, trimming those with incomes under $10 in 2007 dollars. When families broke up, he added inter-household transfers to families receiving them and subtracted them from families paying them. Some of these details are found in Gosselin and Zimmerman (2008).

\(^{17}\) Gosselin and Zimmerman (2008). The authors use the PSID, including the SEO sample, and weight the data. The sample includes adults age 25-64, and families with less than ten dollars in income (in 2007 dollars) before out-transfers are excluded. Gosselin and Zimmerman drop the top and bottom 2% of income changes. The events examined included divorce/separation, death of a spouse, birth of a child, reduced hours due to retirement or disability, head’s unemployment, work loss of head due to illness, and a fall in work hours of the wife. The trends were similar when looking at adults age 35-55 years old and when adjusting incomes for family size. A description of earlier analyses along these lines may be found in Gosselin (2004c).
account for fewer income drops, making up two-thirds of drops in the first period but just half in the second period.

Another recent paper garnering a sizable amount of attention (also based on the PSID) is a Brookings Institution working paper by Karen Dynan, Doug Elmendorf, and Dan Sichel. They find that the risk of a two-year family income decline of 50 percent or more grew unevenly from 6 percent in the early 1970s to 11 percent in the early 2000s, following a notable cyclical pattern. On the other hand, an earlier draft of the paper showed that the share of adults experiencing a 25 percent decline grew from about 5 percent to about 11 percent. Once again, the downward mobility levels in these two drafts are incompatible (due to the percentages being calculated two different ways), though the trends show the same pattern.

A forthcoming paper by Stephen Rose and me uses the PSID and finds that the probability of a 25 percent drop in family income over two years rose and fell cyclically between 1969 and 2004 but increased over the entire period. However, they present evidence that this rise is entirely due to a shift in the early 1990s that is likely an artifact of PSID administrative changes. The Congressional Budget Office recently published their latest analysis of income volatility trends, which constitutes one of the rare studies

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18 Dynan et al. (2008). The authors exclude the SEO and immigrant samples of the PSID. Their sample includes adults who were heads in both years, at least 25 years old and not retired. The authors bottom-code incomes at $1 and adjust them for inflation using the CPI-U. Top codes are applied to cap the same share of the sample in each year. The authors measure percent changes as the two-year change in income, divided by the average of the two years of income. They drop adults in households where the head reports $0 in earnings but over 120 hours of work.

19 Dynan et al. (2007). Methodological details are the same as in their 2008 paper, except that they exclude reports of $0 in earnings by the head are dropped and in computing percent drops, the change in earnings between years $t$ and $t-2$ is divided by the average of years $t-2$, $t-3$, and $t-4$.

20 Rose and Winship (forthcoming). The authors include the SEO sample and apply the weights. The sample consists of adults age 26-59. They do not adjust family income for family size but adjust for inflation using the CPI-U-RS linked to the CPI-U.
that does not rely on the PSID.\textsuperscript{21} CBO researchers (henceforth Dahl et al.) analyzed a restricted-use version of the Survey of Income and Program Participation (SIPP) linked to individual Social Security earnings data. They relied on the Social Security data to measure adults’ earnings and combined them with non-labor income from the SIPP. Analyzed this way, the share of households with one-year income drops of 25\% or more was flat from 1985 to 1993, declined from 1993 to 1998, and then increased from 1998 to 2005. However, the changes were small – with estimates ranging from 11\% in 1985 and 1993, to 9\% in 1998, to 11-12\% in 2005. Peter Orszag presented other CBO results, showing trends in the risk of a 50\% income drop.\textsuperscript{22} This probability declined from about 4.4\% in 1985 to about 3.9\% in 1994 and then to about 3.5\% in 1998; it increased to about 4.4\% again by 2002.

Turning to upward absolute mobility, Gosselin found that the probability of a 50\% increase in income rose from 12 percent in the 1970s to nearly 16 percent in the 1990s, then fell to 14 percent in the early 2000s.\textsuperscript{23} Dynan et al. found an increase from 7 percent in the early 1970s to as much as 13 percent in the late 1990s before upward

\textsuperscript{21}Dahl et al. (2008). The authors use the 1984, 1990-93, 1996, 2001, and 2004 panels of the SIPP data matched to the Detailed Earnings Record of the Social Security Administration. They measure wage and self-employment income from the DER and non-labor income from the SIPP, adjusting for inflation using the CPI-U-RS. The sample includes households headed by an adult age 25-55. Households had to be in the SIPP panel for 24 months. The authors dropped households in the top or bottom 1\% in either year, as well as households with a member age 18-64 in month 12 of the panel who was not successfully matched. SIPP panel members who were not matched had heads with less education and more variable SIPP incomes than those matched, but the bias did not change over time. Self-employment earnings are topcoded prior to 1991, affecting the 1984, 1985, and 1990 income measures, but excluding self-employment earnings did not affect the trends they estimated. Percent changes were measured as the one-year change in income divided by the average of the two years.

\textsuperscript{22}Orszag (2008). The estimates are from SIPP data matched to Social Security Administration data.

\textsuperscript{23}Gosselin (2008).
mobility dropped to 9 percent in the early 2000s.\textsuperscript{24} Dahl et al. found a slight decline in the share of households experiencing a 25% gain, falling from 18 percent in the mid-1980s to about 15 percent in the early 2000s. Orszag reported a bigger decline in the share experiencing a 50% gain—4 percent in 1994 and in 2002 versus 8 percent in 1985. Rose and Winship report little change in the probability of a gain of 25 percent over two years from 1969 to 2004. Hacker has never reported trends in upward mobility.

A number of other studies consider trends in upward or downward absolute income mobility. Greg Acs and his colleagues found that the share of adults experiencing a 25% drop in income between successive four-month waves of the SIPP over the course of a year rose from 40 percent in 1996 to 46 percent in 2001 and then fell to 36 percent in 2004.\textsuperscript{25} The share experiencing a 50% drop rose from 14 to 18 percent and then fell to 15 percent, and the percent experiencing a 75% drop rose from 4.0 to 5.1 to 5.6 percent.

Gottschalk, McLanahan, and Sandefur found that the likelihood of staying poor in two consecutive years was fairly constant from the early 1970s to the mid-1980s, with just under 40 percent escaping poverty throughout most of the period.\textsuperscript{26} Duncan, Smeeding, and Rodgers examine transition matrices based on fixed real income cutoffs.\textsuperscript{27} They find that upward mobility from the “low” income category to the “middle” or

\textsuperscript{24}Dynan et al. (2008). Dynan et al. (2007) found similar patterns for the likelihood of a 25% income gain.

\textsuperscript{25}Acs et al. (forthcoming). The authors use the 1996, 2001, and 2004 panels of the SIPP, focusing on family heads and partners age 25-61 who are in families with children. I thank Greg Acs for sharing these results with me.

\textsuperscript{26}Gottschalk, McLanahan, and Sandefur (1994). The authors use the PSID but otherwise do not elaborate on the methodological details of their study.

\textsuperscript{27}Duncan, Smeeding, and Rodgers (1993). The three categories included low ($0-18,500 in 1987 dollars), middle ($18,500-$55,000), and high. The authors found similar results when they looked at mobility using income-to-needs measures based on poverty thresholds: low (0-2.0), middle (2.0-6.0), and high. For additional methodological details, see footnote 8.
“high” categories was lower during the 1980-87 period than during the 1967-1980 period (30 percent versus 36 percent). Similarly, downward mobility from the “high” category was also lower (27 percent versus 31 percent). In contrast, both upward and downward mobility from the “middle” category increased (by roughly a similar amount).

Tom Hertz, in two recent papers relying on matched interviews from the Current Population Survey, examined trends in several measures of absolute mobility.\textsuperscript{28} In general, absolute downward and upward mobility between two years both increased between 1987 and the early 1990s and again from the early 1990s to 2004. The share of households with declines of 50% or more and with gains of 50% or more both grew from 1987 to 1994 and then increased by a smaller amount between 1994 and 2005. The median gain increased from 1987 to 1994 and again from 1994 to 2005. The median loss increased from 1987 to 1994, but was flat from 1994 to 2005. The median income gained and the median lost both rose steadily from 1991 to 2004.

The share of adults losing $20,000 or more increased from 13 percent to 15 percent between 1991 and 1998 and then again to 17 percent by 2004. The percent gaining $20,000 or more in one year increased from 1991 to 1998 and then was flat. The percent of adults with income gains of any size rose from 51 percent to 55 percent, then fell back to 51 percent again. Overall, downward mobility became increasingly more prevalent compared with upward mobility, as can be seen in the fact that the median

\textsuperscript{28} Hertz (2006, 2007). Hertz provides few methodological details in his 2006 paper. In the 2007 paper, he excludes households where imputations account for more than 10% of income, households that don’t have at least one person in common in both years, households with income in the top 1.5% in either year, and households with incomes below $500 (in 2005 dollars). He compares three points in time in each study, but they are not the same points of time for both studies. In the 2007 paper he examines log income per person, while in the 2006 study he examines untransformed income.
change in log income per person declined from 1987 to 1994 and then declined again between 1994 and 2005.

When Hertz defined quantiles according to fixed 2004 levels, upward absolute mobility from the bottom decile or quintile increased between 1991 and 2004, with the share gaining income, the median income change, and the median increase all growing steadily over the period. Downward mobility from the top decile or quintile increased between 1991 and 1998, but then was flat or declining. In the middle quintile, upward and downward mobility both increased from 1991 to 1998, while downward mobility increased and upward mobility declined between 1998 and 2004.

Finally, several studies allow one to consider trends in non-directional absolute income mobility. Dahl et al. reported that the share of households with one-year income changes (increases or decreases) of 25% or more declined between 1985 and 1991 but changed little thereafter, holding at about 25 percent of households through 2005. Using the SIPP data alone, without linking to the Social Security data, showed the same trend when the authors excluded households with imputed earnings data. Orszag showed that the share experiencing a change of 50% or more declined between 1985 and 1994 and was then flat through 2002.

Dynan et al. showed that the share of household heads experiencing a 25-percent change in income over two years increased from the early 1970s to 2000. It then declined in the early 2000s. In contrast, Rose and Winship find no change over time. Hertz found that the median absolute value of the change in log income per person increased from 1987 to 1994 and increased slightly again from 1994 to 2005.\(^{29}\)

\(^{29}\) Hertz (2007).
median absolute change in household income increased from $8,045 in 1991 to $10,874 in 1998, and then to $11,345 in 2004.\textsuperscript{30}

Austin Nichols and Seth Zimmerman, using the PSID, show results for the trend in the average (across adults) absolute value of the change in income over five-year periods, controlling for income in the middle year.\textsuperscript{31} This measure of absolute mobility increased fairly steadily from the early 1970s to the late 1990s, then declined back to its mid-1990s level.

In sum, there is little consistency in the various estimated trends in absolute income mobility. The one point on which the research agrees is that absolute mobility—non-directional, upward, and downward—increased in the 1970s. The estimates from CBO’s SSA-based data consistently show only small changes in mobility, generally in contrast to the results from the PSID-based studies. But the PSID-based studies often disagree among themselves. Most of the evidence from the PSID indicates an increase in mobility over the 1980s and an increase in downward mobility during the late 1980s. The PSID-based studies all agree that mobility was higher in 2000 than in 1970, and the Dynan et al. and Hacker studies find clear counter-cyclical patterns in directional mobility, in contrast to the findings for relative directional income mobility. Little else can be said comparing the relative and absolute income mobility results.

\textsuperscript{30} Hertz (2006).

\textsuperscript{31} Nichols and Zimmerman (2008). The measure is technically the predicted absolute value of the change in income from a Poisson regression model that includes year dummies and income measure in the middle year. They use family income in the PSID, including the SEO, and include individuals age 25-61. Income is adjusted for inflation using the CPI-U-RS. This measure shows the same trend regardless of whether incomes of $0 are excluded or not.
Comparing absolute income mobility findings to the evidence on absolute earnings mobility trends is also difficult due to inconsistencies. It appears that both men’s earnings and income mobility increased in the 1970s. If the PSID results are to be believed, then income mobility increased during the 1980s and 1990s as well, but that is not the case for earnings mobility (or for income mobility based on the SSA data). The cyclicality of directional earnings mobility is mirrored in the results for directional income mobility.\textsuperscript{32}

**Research Examining the Dispersion of Income Changes**

As with earnings instability, the next logical step after looking at changes in the typical person’s absolute income mobility is to look at changes in the dispersion of absolute mobility. Dynan et al. (2008) looked at the combined earnings of heads and their spouses and tracked the standard deviation of two-year percentage changes.\textsuperscript{33} They found volatility increasing by about 30 percent from the early 1970s to the early 2000s, mostly from the mid-1970s through the early 1980s and during the early 1990s. Similarly, combined head and spouse transfer income volatility rose by about one third, mostly during the early and mid-1970s and early 1990s. On the other hand combined head and spouse capital income volatility only increased by about 10 percent over the

\textsuperscript{32} In addition to these income volatility studies, Davis and Kahn (2008) report that the mean of the absolute value of the logged change in household consumption over six months was larger – indicating more absolute mobility – during the 1992-2004 period than in the 1980-91 period. The authors rely on the Consumer Expenditure Survey and examine consumption expenditures, defined as those on nondurable goods and services, per adult equivalent. Trends are grouped by predicted consumption decile (predicted using the first quarter of consumption data from demographic variables). Volatility increased for the third through tenth deciles. Conley and Glauber (2008) found that short-term absolute wealth mobility rose between the second half of the 1980s and the early 2000s. The authors examine typical changes (in dollars) over five- or two-year periods in the PSID, as well as the share with changes of a given dollar amount.

\textsuperscript{33} Dynan et al. (2008). The authors measure volatility as the standard deviation of percent changes in income over two years, where the percent changes are computed as described in note 18.
period. Finally, total household income volatility increased by about 35 percent from the early 1970s through the early 2000s, with a rapid increase in the early 1990s.

The CBO has also estimated trends in the dispersion of percent changes in income. Orszag reported that volatility dropped between 1985 and 1994, was flat from 1994 to 1998, and then increased slightly from 1998 to 2002.\textsuperscript{34} Hacker and Jacobs reported that the standard deviation of two-year changes in income increased 51 percent from 1971 to 2004.

Nichols and Zimmerman showed that the standard deviation of percent changes in income over four years increased unevenly from the early 1970s to the mid-1990s, was flat through the late 1990s, then declined.\textsuperscript{35} The increase from the early 1970s to the early 2000s was just 5 percent. When families reporting $0 in income and other suspect observations were dropped, the trend was dominated by noise.

Craig Gundersen and James Ziliak also produced PSID-based estimates of income-change dispersion.\textsuperscript{36} They captured the residuals from a regression model predicting one-year income changes from demographic variables, time dummies, and individual fixed effects and then tracked the mean of the squared residuals in each year. They found that volatility fell, rose, and fell again over the 1980s, increased dramatically

\textsuperscript{34} Orszag (2008). No methodological details are given beyond the source of the data, which is the SIPP matched to SSA data.

\textsuperscript{35} Nichols and Zimmerman (2008). They computed percent changes in the same way as Dynan et al. (2008) for illustration.

\textsuperscript{36} Gundersen and Ziliak (2003). The authors exclude those less than age 25 in 1980, those not in the sample for at least three years, students, the disabled, the institutionalized, those with more than a 300% increase or a 75% decrease in income or consumption, those with less than $520 in annual food consumption and those with under $1,000 in annual income. Income includes earnings, transfers, and capital income and excludes federal taxes. The authors model income as function of observable demographics, a time-invariant permanent effect, a random growth term, year fixed effects, and a random walk component. They then estimate the change in income using OLS, capture the residuals, and measure volatility as the average squared residual for each year.
in the early 1990s before plunging, and then increased modestly through 1997. Over the entire period, volatility rose by about 10-15 percent (measured in standard deviations).

Finally, a new paper by Richard Blundell and his colleagues reports trends in the variance of the one-year change in residualized log post-tax income among married couples, using the PSID. They find that dispersion in income changes increased between 1981 and 1987, fell between 1987 and 1991, and increased in 1992, rising over the 1980-1992 period by about 40 percent (or 22 percent if measured in standard deviations).

These few studies on dispersion in income changes agree that income volatility increased in the 1970s and changed little in the 1980s. The research based on the PSID finds increases in volatility in the 1990s, though CBO’s declining trend from the mid-1980s to the early 2000s is inconsistent with the PSID results. Finally, the trend in volatility in the early 2000s is inconsistent across studies. These general findings are consistent with the analogous literatures on trends in absolute income mobility and trends in male earnings dispersion in that all find increases in volatility in the 1970s and all find higher volatility in the early 2000s than in 1970 or 1980 in the PSID-based studies. On the other hand, the CBO research implies declining volatility for both earnings and income since the mid-1980s.

37 Blundell, Pistaferri, and Preston (2008). The authors use the 1978-1993 waves of the PSID, excluding the SEO. The sample consists of continuously married couples headed by a male, and the income measure is labor income plus transfers, minus the estimated federal taxes on these components. Federal taxes are imputed in 1992 and 1993 using the NBER TAXSIM program. The authors exclude those with topcoded income, financial income, or federal taxes, with growth of 500% or more or -80% or less, or with <$100 of income. They also exclude those with missing data on education or region, born before 1920 or after 1959, or less than 30 or over 65. Log income is residualized by regressing it on year dummies, year of birth dummies, and family characteristics and capturing the residuals.

Research Summarizing Within-Person Income Dispersion Across Years

In Appendix One, I showed that previous research on earnings volatility measured as the typical within-person dispersion over a series of years was unclear as to whether volatility had increased since 1980. I now examine the literature on within-person dispersion in family income. To review, there are three basic measures that have been used in earnings volatility research. Peter Gottschalk and Robert Moffitt pioneered the first two types of measure.39 The first defines volatility as the average within-person variance of incomes over some period, comparing two fixed and non-overlapping periods. The second is similar except that it looks at the mean within-person variance within some window centered on particular year and allows the window to move as different years are considered. The third measure, used by Gosselin, defines volatility as the typical maximum absolute change in income over some moving window.

Hacker reported income volatility results for a fourth measure – a person’s lowest family income over some period expressed as a percentage of his or her highest income over the period.40 His revised results indicated that this figure fell from 42 percent for the 1971-1980 period to 37 percent for the 1981-1990 period, to 28 percent for the 1991-2004 period, indicating rising volatility.

Gottschalk and Moffitt presented income volatility results using one of these methods only once, at a workshop organized by The Pew Charitable Trusts’ Economic

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40 Hacker (2008). He dropped all non-positive incomes, then he trimmed the top and bottom 2% of the remaining observations. See note 10 for additional methodological details. Hacker’s results in the first edition of GRS indicated that the decline was from 43% to 36% to 25% from 1973-82 to 1983-92 to 1993-2002.
They showed that income volatility measured using the PSID as the mean of the within-person variance of income over a nine-year window was flat from 1974 to 1980, then steadily rose through 1998, accelerating in the late 1980s. Over the entire period, volatility increased about 80 percent (which would be less if the results were presented as standard deviations).

Benjamin Keys used Gottschalk’s and Moffitt’s first method, also using the PSID. He found that income volatility was higher in the 1990s than in the 1980s, which featured higher volatility than the 1970s. The pattern was true for men and women, blacks and whites, ranging from a 46-percent increase among white women from the 1970s to the 1990s to a 228-percent increase among black men.

Gosselin and Seth Zimmerman used the PSID to measure volatility as the variance of family income over four years in a seven-year window. They found that mean volatility rose between 1970 and 1980, flattened out in the early 1980s, rose at an accelerating pace from 1983 to 1992 and then flattened again through 1998. It doubled over the entire period. Median volatility was flat from 1970 to 1985 before increasing through 1998, rising about 75 percent over the entire period. Furthermore, when Gosselin and Zimmerman adjusted family income for needs by dividing by the poverty line, the results were basically unchanged. Finally, the authors checked their results

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41 Gottschalk and Moffitt (2007). The authors use the PSID, and examine log family income adjusted for needs by dividing by the poverty line for a family of a given size. They trim the bottom two percent of income.

42 Keys (2008). The author focused on non-student heads age 20-59. He dropped $0 observations, trimmed the top and bottom 1 percent of incomes, and regressed logged family income on a quartic in age. Volatility is estimated within-person over 10 years.

43 Gosselin and Zimmerman (2008). The authors first regress log family income on a quartic in age, then compute various quantiles of the variance of the residual over years $t$, $t+2$, $t+4$, and $t+6$. Individuals must have at least three valid observations in these four years to be included. For additional methodological details, see note 17. Few details are provided for the SIPP analyses, but volatility is measured as the mean within-person variance of income over three consecutive years.
using the SIPP. They found that volatility was flat from 1983 to 1993 and rose from 1993 to 2001, an increase of roughly 50 percent over the entire period. Comparing years in which volatility estimates are available for both the SIPP and the PSID, Gosselin’s and Zimmerman’s estimates indicate that volatility rose about 25 percent in the SIPP from 1983 to 1996 and by about two-thirds in the PSID.

Nichols and Zimmerman looked at trends in the typical within-person standard deviation of family income over a five-year window using the PSID. Mean volatility of income levels rose steadily from the early 1970s through the 1990s and then fell between 2000 and 2004. Mean volatility of logs rose slowly from the early 1970s to the mid 1980s, was flat through the early 1990s, rose through 2000 and then fell. Over the whole period volatility of levels rose about two-thirds while log volatility rose about one-third. Median volatility, however, increased by about half these amounts. Median volatility of income levels fluctuated shallowly between the early 1970s and the early 1990s, then rose through 2000 and declined through 2004. Median log income volatility showed essentially the same pattern. The magnitude of the increase in volatility over time depends not just on whether levels or logs, or means or medians are examined but on whether and how low incomes are trimmed.

Yet another PSID-based analysis of within-person dispersion trends comes from Gosselin’s High Wire. Gosselin measured volatility as the 68th percentile across adults

44 Nichols and Zimmerman (2008). Within-person dispersion is measured over years $t-2$, $t$, and $t+2$, and only those with valid incomes in all three years are included. The authors also show trends when other transformations of income are used, including the cube root, the inverse hyperbolic sine, and the neglog transformation. They generally align with the logged results. The authors also report that the results are similar if they adjust family income for needs. For additional methodological details, see note 30.

45 Gosselin (2008). Gosselin computed, for each individual, the maximum absolute value of the annual change over a five-year window (including years $t$, $t+1$, ..., $t+4$). He then reported the 68th percentile across all adults, or across all adults within some quantile (based on five-year average income). Individuals
of the maximum real income change (in absolute value) experienced over three years within a five-year window. He found that volatility increased from the early 1970s through the early 2000s, accelerating in the 1990s before leveling off in the early 2000s. He also reported increases in the 68th percentile of the maximum swing for the bottom decile, the middle quintile, and the top decile, with the increase larger at the bottom than at the middle or top.

A rare non-PSID analysis examining within-person income dispersion is a new working paper by Nichols and Melissa Favreault. Using the SIPP matched to Social Security data, and measuring volatility as the variance of combined head and partner earnings over the previous five years, they find that mean volatility declined between 1981 and 1982, increased between 1982 and 2001, and then declined through 2003. For five of six SIPP panels, mean volatility increased about 75 percent over the entire period (which would be lower if expressed in standard deviations).46

In a recent book chapter, Craig Gunderson and James P. Ziliak rely on Gottschalk’s and Moffitt’s original approach (variance of transitory income over non-moving windows).47 Once again using the PSID, they measure transitory variance as the

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46 Nichols and Favreault (2009). The authors use the 1990-93, 1996, and 2004 panels of the SIPP matched to Social Security Administration earnings data. Their sample includes adults born between 1936 and 1956, who were 24 to 44 years old in 1980 and 48 to 68 years old in 2004. They drop the largest one percent of variances.

47 Gunderson and Ziliak (2008). The authors use the 1980-2003 waves of the PSID. They exclude households where the head was in the sample less than three years, was under 25 in 1980, was a student,
within-person variance of deviations from average income over a five-year period. They find declines in mean volatility for five age groups between 1980-84 and 1985-89 followed by increases between 1985-89 and 1990-94. From 1990-94 to 1995-2002, volatility declined among those under age 30 or older than 50, was flat among adults in their thirties, and increased among those in their forties. Overall, volatility declined in this last period. Over the entire period covered by the paper, volatility was mostly unchanged, but it rose among people in their thirties.

Finally, two studies measure volatility as the within-person coefficient of variation over a fixed number of years. Using the PSID, Lily Batchelder regressed volatility of log taxable income over a six-year period on a number of variables, including a linear time trend that broke the years 1968 to 1992 into four equal time periods (i.e., a trend that imposed the restriction that the change between each period be the same magnitude).\(^48\) She found an increase in volatility over the twenty-four years.

Neil Bania and Laura Leete used the SIPP and measured volatility as the within-person coefficient of variation of monthly income over the previous twelve months.\(^49\) They found that mean volatility of family income in a sample that was relatively low income increased by 17 percent between (roughly) 1992 and 2003. Mean family earnings volatility was flat, while the volatility of AFDC/TANF and of other income was flat.

\(^{48}\) Batchelder (2003). She includes the SEO and weights the data. Her sample includes families with heads age 44-49. Batchelder bottom codes incomes at $0 and recodes all top-coded incomes to the mean taxable income of U.S. taxpayers earning above the top code in a given year.

\(^{49}\) Bania and Leete (2007). The authors use the 1991, 1992, and 2001 SIPP panels, covering October 1991 to December 1992 for the 1991 and 1992 panels and October 2002 to December 2003 for the 2001 panel. Their sample includes household heads age 18-59 who were in households of at least two people and with non-negative pretax incomes no more than 300% of the poverty line.
When the value of food stamps and WIC benefits were added to family income, volatility rose by 19 percent.

These within-person dispersion studies, taken as a whole, show a secular rise in mean volatility from the early 1970s to the early 2000s, with the increase (expressed as variances) ranging from roughly two-thirds to 100 percent. Median volatility increased less dramatically according to Nichols and Zimmerman, but Gosselin and Zimmerman report a sizeable increase (though because median volatility levels are relatively low, there is presumably a considerable amount of imprecision in their estimate of the rise over time). While not entirely consistent, the studies generally find volatility increasing in the 1970s, 1980s, and 1990s, but flat or declining in the early 2000s. This rise is much clearer than the probable increase in within-person earnings dispersion among men found in the Chapter Two, and it is more consistent across time than the increase in income-change dispersion.\(^{50}\)

**Research Summarizing Across-Person Dispersion of Income Shocks**

My review of the existing literature on income instability and volatility concludes by examining the research organized around “transitory variances”. Briefly, the idea behind this approach is to elaborate formal models of individual earnings dynamics, with earnings typically disaggregated into a permanent and transitory component. The models specified imply restrictions on the variances and covariances of earnings across years. In

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\(^{50}\) In addition to these studies, Olga Gorbachev (2008) found an increase in food consumption volatility of 1.0% annually from 1974-78 to 1998-2002, with the increase coming almost entirely in the 1996-2002 period. Keys (2008) also found increases in food consumption volatility in the 1970s, 1980s, and 1990s (except among white women in the 1990s).
the more sophisticated models, the parameters of the model are estimated by finding the set that produces the covariance matrix that best fits the observed covariance matrix in the data. The parameters of simpler models (with stronger assumptions) can often be estimated without such complicated techniques, as with the variance decomposition model of Peter Gottschalk and Robert Moffitt that Hacker has relied on.

It is unclear that models developed to describe earnings dynamics of individuals are appropriate for modeling income dynamics of households. Households include multiple potential earners who can change their labor force participation and work decisions depending on the decisions of others. Furthermore, income sources other than earnings—such as investment income or public transfers—may not follow the same dynamics as earnings.

For example, in the variance decomposition model of Gottschalk and Moffitt, as discussed in Chapter Two, earnings are assumed to be fixed over a window of years, save for an annual random shock that rapidly dissipates. Income volatility research using the variance decomposition model must assume that all sources of income in a household are subject to the same random shock in every year—that all sources of income are fixed save for a single annual shock that affects one or more of them. For this reason, far fewer studies of income volatility have been conducted using this approach than studies of earnings volatility.

Using the variance decomposition model, Hacker’s pre-tax family income volatility estimates, as reported in *GRS*, indicate a fairly steady increase in volatility from
1974 to 1984, with local peaks that tend to lag recessions by a year or two.\textsuperscript{51} Volatility declined slightly between 1984 and 1989 and then took off between 1989 and 1994. It then declined just as sharply between 1994 and 1998, with a small increase in 1996 interrupting the decline. Volatility spiked upwards again from 2000 to 2002. In all, volatility increased by 360 percent from 1974 to 1994, declined by 57 percent from 1994 to 1998, and increased 50 percent from 2000 to 2002. Volatility more than doubled from 1991 to 1993 alone according to these results. Post-tax income volatility showed similar patterns and grew 130 percent over the entire 1974-2002 period.

The next iteration of results from Hacker arrived in the “Revised and Expanded Edition” of \textit{Great Risk Shift}, published in early 2008.\textsuperscript{52} This time, rather than volatility increasing 375 percent between the early 1970s and its 1990s peak, the rise Hacker showed was one-third that. The new pre-1990s trend was noisier than Hacker’s original estimates, with a sizable decline in volatility during most of the 1980s after the run-up in the first years of the decade. The trend from the early 1990s onward showed a large increase in volatility followed by a large decline in the mid-1990s, with another notable increase from 1998 to 2002. From 1973 to 2004, Hacker showed an increase in volatility of about 95 percent (presented as variances).

\textsuperscript{51} Hacker (2006). In computing covariance terms, Hacker used 5-year lags, though to obtain the 2002 estimate, he used a 6-year lag, since there was no 1998 survey to provide 1997 income estimates. For additional methodological details, see note 14.

\textsuperscript{52} Hacker (2008). He dropped all non-positive incomes, then he trimmed the top and bottom 1% of the remaining observations. In computing covariance terms, Hacker used 4-year lags. For additional details, see note 10.
Just months after the revised edition of *GRS*, Hacker and Jacobs produced yet another set of estimates, for an Economic Policy Institute research brief.\(^5\) In this most recent version of Hacker’s figures, volatility is shown rising 150 percent from 1973 to 1993, falling by more than half between 1993 and 1998, then increasing between 2000 and 2002. Over the entire 1973-2004 period, volatility essentially doubled. The early-1980s increase in volatility is not as sharp as in the revised edition of *GRS*, but the latest figures still imply that volatility increased by something like 75 percent just between 1991 and 1993.

These results followed the earlier estimates of Hacker but also previous estimates by Jacobs.\(^5\) Jacobs found a bigger increase in the transitory variance of income between the early 1970s and the early 1980s than in her subsequent paper with Hacker. She omitted the 1991 to 1995 data points in her chart, owing to a lack of trust in the estimates given the data issues I review in Chapter Three.\(^5\) Her estimates for 1996 to 2004 were similar to her later results, but shifted downward. In all, Jacobs reported a roughly 70 percent increase in income volatility from 1973 to 2004 (expressed in variances).

Gottschalk and Moffitt produced trends in the transitory variance of income only once, in their presentation to The Pew Charitable Trusts/Brookings Institution workshop,

\(^{53}\) Hacker and Jacobs (2008). Once again, Hacker’s analyses rely on the PSID’s pre-tax family income variable. Incomes are again adjusted for family size by dividing by the square root of family size, and they are once again logged and adjusted for inflation using the CPI-U. Again he looks at individuals age 25-61, including the entire core sample and using the sample weights. He dropped all incomes of $1 or less, then trimmed the top and bottom 1% of the remaining observations. Once again, four-year lags are used in computing covariance terms.

\(^{54}\) Jacobs (2007). Jacobs used the PSID and focused on persons age 25 to 61. She logged household income and trimmed the top and bottom 3%. Business owners were excluded from the sample. She used a four-year lag for the covariances in computing permanent variances.

\(^{55}\) Though unacknowledged in Jacobs’s paper, she and I collaborated extensively in the initial stages of her research.
and they made no effort to publish those results.\textsuperscript{56} They found that the transitory variance declined modestly over the early 1970s, then increased slowly through the mid-1980s. Volatility then declined slightly over the rest of the decade before increasing notably in the early 1990s. Volatility fell between 1993 and 1998 and then increased again through 2002. Over the entire period, the transitory variance increased by roughly 40 percent (or about 20 percent expressed in standard deviations).

Nichols and Zimmerman also presented PSID-based trends in volatility using the Gottschalk-Moffitt variance decomposition method.\textsuperscript{57} They found that volatility increased from the early 1970s to the early 1980s, then declined through the mid-1980s. It rose in the late 1980s and early 1990s, before falling again through the mid-1990s. Finally, volatility rose again in the late 1990s and early 2000s. The increase over the entire period was 20 to 48 percent (in terms of standard deviations) depending on how they trimmed high and low incomes.

Rather than this variance decomposition model, two studies by Richard Blundell and his colleagues use more sophisticated error components modeling to generate estimates of transitory income variance. Blundell and his colleagues use a model that includes a permanent household income component following a random walk and a transitory component modeled as a moving average process.\textsuperscript{58} In this model, there are

\textsuperscript{56} Gottschalk and Moffitt (2007). The authors use the PSID and the variance decomposition model that defines the transitory variance as the difference between the total variance and the covariance between incomes measured in two years. For additional methodological details see note 40.

\textsuperscript{57} Nichols and Zimmerman (2008). The authors used a four-year lag to compute the covariances. For additional methodological details, see note 30. The authors also present results using income levels instead of logs, however the Gottschalk-Moffitt variance decomposition model applies only to logged income (see Chapter Three above).

\textsuperscript{58} Blundell, Pistaferri, and Preston (2008). The authors use the PSID, excluding the SEO sample. They focus on households headed by a male age 30 to 65 who was continuously married to his wife. Income is
both permanent and transitory shocks. The authors found that the variance of transitory shocks declined in the early 1980s, increased notably through 1986, declined slightly through 1989, and increased slightly through 1992. The variance ended 25 to 50 percent higher than in 1980, with the increase occurring entirely in the mid-1980s. The variance of permanent shocks increased in the early 1980s, flattened out in the mid-1980s, then declined through 1992, ending perhaps 35 to 50 percent higher than in 1980.

Blundell and Luigi Pistaferri estimate a similar model using the PSID, which models both income and consumption dynamics as depending on shocks to permanent income. They find that for both low-income and medium- to upper-income households, the variance of the permanent income shock is smaller in 1985-92 than in 1979-84. The variance of the transitory component – which the authors assume represents measurement error – follows the same pattern. When they instead assume that the transitory component is a shock to income, the changes in the variance of the shocks estimated by their model become statistically insignificant.

These studies all use the PSID and are generally consistent. Volatility increased in the 1970s (except in Gottschalk and Moffitt), in the 1980s, in the 1990s (except in measured as household labor income plus transfers, minus estimated federal taxes on these components. Federal taxes are imputed in 1992 and 1993 using the NBER TAXSIM program. The authors exclude households with top-coded income, financial income, or federal taxes; with year-to-year growth of 500% or more or -80% or less; with less than $100 of reported annual income; with missing data on education or region; or born before 1920 or after 1959. Prior to estimating their model, the authors regress log income on year dummies, year of birth dummies, and various demographic characteristics.

59 Blundell and Pistaferri (2003). Income is measured as household labor income plus cash transfers, adjusted for inflation using the CPI-U. The sample is restricted to continuously married couples headed by a man, with no family composition changes. They exclude those younger than age 25 or older than 65, those with top-coded income in any year or with missing data for education or region or who are income outliers. The authors’ model derives from a utility-maximization framework. Income is modeled as having a permanent component following a random walk and a transitory component. The permanent shock also affects year-to-year changes in consumption, which otherwise follow a constant path over time conditional on demographic controls. They categorize people as low-income if their pre-1979 average income was below 200% of the federal poverty line.
Nichols and Zimmerman), and from 1998 to 2002. These results are consistent with the trends in within-person income dispersion and in the dispersion of earnings shocks noted in the Chapter Two. The increase from the early 1970s to the early 2000s was somewhere between 20 and 50 percent expressed as standard deviations, according to these studies. That is a much smaller increase than reported by Hacker and Jacobs, who summarize the previous research in a graph showing increases in the range of 50 to 120 percent (presented as *variances*).\(^6^0\) It is comparable to the range found in the literature on male earnings volatility that use the same methods, summarized in Chapter Two as running from 15 to 65 percent.\(^6^1\)

**Other Approaches to Volatility Measurement**

Austin Nichols decomposed variation in incomes across people and years into inequality, volatility, and mobility.\(^6^2\) He measured inequality as between-person variation in mean incomes measured over five years in the previous nine-year period (using income measured every two years). Within-person variation in demeaned income is then divided into mobility and volatility, where mobility is measured as within-person, across-year variation in predicted income (predicted from a linear time trend) and volatility is measured as within-person variation in individual incomes around individual trends.

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\(^{60}\) See [http://pantheon.yale.edu/~jhacker/volatility_graph.pdf](http://pantheon.yale.edu/~jhacker/volatility_graph.pdf).

\(^{61}\) In addition to these studies, Gorbachev (2008) finds an increase in consumption volatility using a model that considers both income and consumption dynamics.

\(^{62}\) Nichols (2008). Nichols uses the PSID, and his income measure consists of family income plus food stamps. He estimates results with and without taxes, imputed using NBER’s TAXSIM. He also estimates them with and without adjusting family size, by dividing by the square root of family size. Nichols drops the top 2% of incomes in every year. The sample includes adults age 30 to 60.
Nichols found that volatility increased from 1976 to 2004, increasing more sharply from 1984 to 1990 and from 1998 to 2004. When incomes were adjusted for family size, volatility was relatively flat from 1976 to 1985, rose sharply through 1988, changed little through 1998, then increased. Over the entire period the increase was about 40 to 60 percent. Mobility fell from 1976 to 1981, rose sharply from 1981 to 1986, then showed little consistent trend. When adjusted for family size, mobility followed a similar pattern but declined from 1988 or so until about 1996 before increasing. Over the entire period, the increase was about 30 to 50 percent. As the window of time used to examine inequality and mobility is widened, the rise in volatility is sharper and the increase in mobility falls. When Nichols used logs instead of levels, the volatility increase is more constant over time, and the mobility trends become more exaggerated.

Esfandiar Maasoumi and Mark Trede measure short-term mobility by comparing the inequality of incomes measured over several years to an average of one-year inequalities within the multi-year window.\(^\text{63}\) They found that post-government income mobility among all persons declined from 1985 to 86, increased, then declined in the late 1980s (declining over the period as a whole). When they divided the sample into three age groups (26-35, 36-50, 51-65), the results were inconsistent, except that mobility declined over the entire period for each.

Finally, Richard V. Burkhauser and John G. Poupore, in a paper on inequality trends in the United States and Germany, indicated in a note that income mobility measured as the percentage reduction in inequality when five years of income are

\(^{63}\) Maasoumi and Trede (2002). The authors use the PSID and measure post-government income as pre-government income (which includes the imputed rental value of owner-occupied housing) plus public transfers, minus income and payroll taxes. They exclude those with post-government income of less than $450.
averaged versus the average inequality of the five individual years was “slightly” higher in the 1970s than in the 1980s.\textsuperscript{64}

\textbf{Summary of Previous Research on Income Instability and Volatility}

The research on income instability and volatility trends is much more consistent than that on trends for earnings. This fact is largely due to the dominance of PSID-based studies in the research on income instability and volatility trends, and the relatively small number of studies thus far conducted. The research generally finds that short-term relative income mobility was largely unchanged or declined during the 1970s and the 1980s. However, research on absolute mobility and on income dispersion show that income instability and volatility increased in the 1970s, 1980s, and 1990s. There is conflicting evidence for the early 2000s, with research on dispersion of transitory shocks showing an increase and that on within-person income dispersion showing a flat or declining trend.

The big exception to these conclusions is the CBO research, which uses the SIPP linked to Social Security records to correct for measurement error in labor income. The CBO research indicates that income instability was basically flat from the mid-1980s forward. Otherwise, the increases in income volatility since 1970, 1980, or 1990 are consistent with the increases in male earnings volatility over these same periods in the research discussed in Chapter Two. Where quantitative estimates of volatility can be compared between the income and male earnings literatures (studies using dispersion-based measures) the two show increases of similar magnitudes.

\textsuperscript{64} Burkhauser and Poupore (1997). This study also used the PSID.
In short, while the rise in income volatility has not been nearly as large as Hacker’s initial estimates implied, the evidence from the PSID does support his Great Risk Shift hypothesis in that it generally finds steadily increasing income volatility and instability over the past thirty-five years. Some dispersion-based estimates imply a steady rise in volatility of as much as 50 percent over the entire period while the CBO figures suggest that it perhaps did not change at all after the early 1980s. This range is uniformly much lower than the doubling of volatility that Hacker has touted, but it implies potentially high volatility nonetheless. Chapter Three attempts to make more sense of the size and timing of any rise in income instability.
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