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Understanding Declining Fluidity in the U.S. Labor Market[☆]

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Abstract

We document a clear downward trend in labor market fluidity that is common across a variety of measures of worker and job turnover. This trend dates to at least the early 1980s if not somewhat earlier. Next we pull together evidence on a variety of hypotheses that might explain this downward trend. It is only partly related to population demographics and is not due to the secular shift in industrial composition. Moreover, the decline in labor market fluidity seems unlikely to have been caused by an improvement in worker-firm matching, the formalization of hiring practices, or an increase in land use regulation or other regulations. Plausible avenues for further exploration include changes in the worker-firm relationship, particularly with regard to compensation adjustment; changes in firm characteristics such as firm size and age; and a decline in social trust, which may have increased the cost of job search or made both parties in the hiring process more risk averse.

Keywords: Labor market transitions, job turnover, hires and separations, labor reallocation, job creation and destruction, labor market churn, internal migration, demographic trends.
JEL: classification J60, J11, J21, J30, R23.

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Introduction

There is mounting evidence that the U.S. labor market has experienced marked declines fluidity along a variety of dimensions. Examples include the rate of job-to-job transitions (Bjelland et al. [8], Molloy et al. [55]), formation of new firms (Davis and Haltiwanger [19]), hires and separations (Hyatt and Spletzer [44]), and geographic movement across labor markets (Kaplan and Schulhofer-Wohl [47], Molloy et al. [55]). This emerging consensus around a general set of concurrent trends raises obvious questions of whether these trends are related and what is causing them. Moreover, the implications of these trends for the performance of the aggregate economy could be substantial. On one hand, the declines in labor market fluidity could signal a rise in the costs of making labor market transitions, which are likely to have negative effects on aggregate productivity and economic performance. On the other hand, lower labor market fluidity could be a sign that there is less need to make such transitions, in which case the implication for aggregate economic performance may well be positive.

Although a number of researchers working in this area have begun to acknowledge that declines in fluidity can be found across a broad set of measures, most relevant papers to date focus on just one or a few such measures. Consequently, the first goal of this paper is to make progress on understanding whether the trends in various measures of labor market fluidity are related and to determine more concretely when declines in fluidity began. As far as we know, this paper presents the first attempt to bring evidence on a large number of series together in order to determine whether they represent a general underlying trend in the labor market.

Because we seek to date the beginning of the decline in fluidity, we need to examine patterns in labor market fluidity that extend for a lengthy period of time – over the past several decades at a minimum.

Consequently, our analysis focuses on data series that are available since at least the early 1980s. We combine information on labor market flows as measured from the perspective of workers (transitions into and out of employment and job-to-job transitions), flows as measured from the perspective of firms (job creation and job destruction), and flows as measured using interstate migration, a transition that is frequently associated with a job change or a change in labor force participation. Bringing together evidence from a variety of sources and methods of measurement is helpful because it reduces the influence of factors that might be idiosyncratic to a particular measure of fluidity and also smooths measurement error that might be specific to a particular data source.

We begin with an analysis of trends in aggregate data. Using time series techniques, we isolate the low-frequency movements in each of eight labor market flows. Most of these trends present a clear downward trajectory. Moreover, these downward trends appear to be related: using principal component analysis we identify a single factor that explains a large portion of the variation of the low-frequency components of these series and puts a positive weight on all of them. This common component appears to have begun its downward trajectory at least in the 1980s, and possibly earlier. Thus, the downward trend in fluidity predates the early 1990s, adding an important caveat to analyses of the trends in fluidity that are based on data sources that are only available for the past few decades. Using our unified measure, labor market fluidity decreased by 10 to 15 percent over the period that we study. However, this single measure smooths changes across several separate measures of fluidity, which individually decline by as much as one third. A decline of this magnitude implies a marked change in the labor market and suggests that the effects the trend in fluidity could be substantial.

Having established that the decline in labor market fluidity is sizeable and appears to be a phenomenon that has been ongoing for three to four decades, we next turn to the question of why. This analysis is composed of three main sections. First, we examine the role of population demographics to see if changes in the distribution of worker characteristics can explain the declines or if the downward trends are concentrated among certain types of workers. Previous work has shown that shifts in the age distribution of the population, as well as other characteristics of workers—including healthcare-related job lock among those covered by employer-provided insurance—do not explain a substantial portion of the decline in various measures of fluidity (Kaplan and Schulhofer-Wohl [47], Molloy et al. [55]). Other work has ruled out a compositional role for some firm characteristics, (Decker et al. [21], Hyatt and Spletzer [44]). Similarly, we find that changes in the distribution of age, sex and marital status explain no more than half of the trends in labor market flows as measured from the worker perspective. Trends in transitions into and out of employment appear to mirror trends in labor force participation as demographic groups with a secular increase in labor force participation (such as prime-age women) have experienced larger declines in transitions out of employment and increases in transitions into employment, while the reverse is true for demographic groups with a secular decrease in labor force participation. Meanwhile, trends in job-to-job flows and interstate migration are similar for most demographic groups. Putting it all together, although demographics go some way toward explaining some labor market flows, they do not seem to account for the bulk of the decline in transitions that is common across all measures. Thus, explanations for the general downward trend should therefore apply to a wide range of workers.

Our second method of narrowing down explanations is to examine state-level trends in labor market fluidity. Local labor markets vary along many dimensions, so it seems natural to expect whatever is causing the aggregate decline in fluidity to have a larger influence in some locations relative to others.

Following a strategy that is similar to that which we used for the aggregate data, we create a measure of general decline in labor market fluidity for each state that is based on both worker and job reallocation. Although labor market fluidity has decreased in all states, it has fallen much more in some states than others. There is a clear geographic pattern, in that fluidity has declined more in the Mountain and Pacific Census Divisions than in other locations. Surprisingly, this geographic pattern persists even after we control for a wide array of state characteristics, indicating that it is not related to the standard observable attributes of the population or to the industrial composition of firms. It is not obvious to us what might be driving this result, but we think it is worth exploring in future work.

Another outcome of the state-level analysis is that states with a larger share of workers in administrative support occupations and machine operators in the late 1970s experienced *smaller* subsequent declines in labor market fluidity. Workers in these occupations were particularly hard-hit by the secular decline in demand for workers who perform routine-intensive tasks, and so their labor market transition rates may have been boosted as they left old jobs and searched for new ones. Thus, these results suggest that the decrease in fluidity would have been larger absent the secular decline in demand for middle-skilled workers. Moreover, it seems unlikely that the secular change in demand for skill and the accompanying widening of wage inequality could have caused the decline in labor market fluidity. Rather, our analysis suggests the two trends are distinct.

Finally, we directly assess a variety of specific theories by weighing the evidence for and against them. Some of this evidence is assembled from existing research, but we also conduct new analysis to generate additional evidence. As we consider these hypotheses, we find it helpful to divide them into two general categories: those that have benign implications for general economic activity, and those with less benign implications. The benign explanations are ones that imply a reduced need for reallocation, such as reasons for improved worker-firm matches. The less benign explanations generally involve an increase in some cost that has caused labor market transitions to become more difficult.

Regarding the benign explanations, one hypothesis is that the match quality between workers and firms has improved. This trend would likely result in either larger returns to staying in the firm (i.e. larger returns to employer tenure) or higher wages in the initial match. Using three cohorts from the National Longitudinal Survey (NLS), we show that after controlling for returns to industry and occupation tenure, returns to employer tenure are small and have not changed noticeably from the late 1960s to the late 2000s. We also examine long-run trends in starting wages in the NLS and Panel Study of Income Dynamics, and find no evidence of a secular increase in match quality as reflected in higher initial wages. Consequently, it seems unlikely that the decline in labor market fluidity can be explained by better matching between workers and firms. A related hypothesis that could explain less labor market fluidity is that if workers and firms have been investing more in job-specific training, because since this type of investment is associated with reduced separations from employers (Cairo and Cajner [12]). Research on the long-run trends in this type of firm-specific training has found mixed results, and more studies on this topic would be helpful. Finally, a decrease in worker turnover might result from a greater ability of compensation to adjust to changes in the productivity of the worker-firm match. Again, evidence supporting this theory is rather mixed, but further investigation, particularly using matched employer-employee data, seems worthwhile.

Turning to less benign explanations, we consider a number of factors that may have caused changes in the labor market to become more costly: a general decrease in the liquidity of the labor market resulting from a reduction of young workers, a decrease in job search or willingness to take new jobs arising from decreases in social capital, a formalization of the hiring process that makes hiring and firing workers more costly, and increased regulations that inhibit labor market transitions. We find little support for any of these hypotheses with the exception of the social capital channel, where we find evidence weakly suggestive of a role for declining trust. In particular, states with larger declines in the fraction of people who think that strangers are trustworthy have also experienced larger declines in labor market fluidity. This correlation is provocative and more work is necessary to explore the mechanism.

In the final portion of our analysis, we discuss potential implications of the decline in labor market transitions. With fewer workers making labor market transitions, we might expect firms and workers to renegotiate wages less frequently. We find that in the 1980s and 1990s wages were most strongly correlated with the best labor market conditions since the worker-employer relationship began, suggesting that wages were renegotiated when outside labor market conditions improved. In the 2000s wages have become more closely tied to conditions in the worker's first year of employment. Thus, workers appear to be renegotiating wages less frequently.

1. Time Series Analysis

The goals of this section are (i) to identify long-run trends in various measures of labor market fluidity; (ii) determine whether these trends are related; and finally (iii) to determine when declines in fluidity began. To do this, we identify eight aggregate time series on flows in the labor market and use time series techniques to estimate low-frequency trends in each of these series. We then assess the comovement of these low frequency trends and discuss what these trends suggest about the magnitude and timing of declining fluidity.

Labor market flows can be measured from the perspective of workers making a transition or from the perspective of firms changing the number or composition of their employees. For example, new employees at a firm must consist of workers who were formerly unemployed (UE), out of the labor force (NE), or working for another firm (JtJ). Similarly, workers flow out of a firm by transitioning to unemployment (EU), leaving the labor force (EN), or leaving to work for a different firm. These worker flows are sometimes grouped into "hires" and "separations":

$$Hires = NE + UE + JtJ \tag{1}$$

$$Separations = EU + EN + JtJ \tag{2}$$

These transitions are gross flows in that someone moving directly from one firm to another will be counted both as a separation (from their old firm) and a hire (to their new firm). Meanwhile, job flows (from firms' perspectives) are usually measured as a net flow. Specifically, job creation is usually defined as net new jobs in new firms and expanding firms, while job destruction is usually defined as net job loss from contracting firms and firms that have shut down. Notably, the sum of aggregate job creation and job destruction is much lower than the sum of aggregate hires and separations because JtJ transitions are a substantial component of worker flows (Davis and Haltiwanger [19], Hyatt and Spletzer [44], Fallick and Fleischman [26]).

In our analysis, we consider flows as measured from the worker and firm perspectives simultaneously because both sets of variables are measured with error and are subject to idiosyncratic influences that are unrelated to the secular decline in fluidity. By combining them, we think we are more likely to identify a trend that accurately reflects general changes in labor market fluidity.

We start our analysis by considering EU, UE, NE and EN because these four flows are available at a quarterly frequency over a span of more than 40 years. Relative to annual data, the quarterly frequency makes it much easier to isolate business cycle fluctuations from those located at lower than business cycle frequencies. The long time series is essential for determining the nature of low-frequency movements began to turn down. Following the analysis of these four quarterly series, we extend the analysis to include job-to-job flows in order to complete the picture of reallocation from the worker's perspective. Doing so requires switching to an annual frequency and considering a shorter time period. Finally we add in three additional annual series: job creation (JC), job destruction (JD), and interstate migration (IM). Although interstate migration does not directly follow from the identities described above, more than half of all interstate migrants report having moved for a reason related to the labor market. Also, because we measure interstate migration using a separate data source from the other worker flows, we think that including this measure helps to mitigate concerns that the measured declines in fluidity owe to mismeasurement in a particular data source.

In our time-series analysis we adopt a two-step procedure. First, following Stock and Watson [64] we estimate the (smooth) low frequency movement of the series estimating a local mean using a biweight kernel. As Stock and Watson [64] point out, the local means estimated using the biweight kernel are approximately the same as those computed as the average of the (stationary) series over a centered moving window, except that the biweight filter means are less noisy because they avoid the sharp cutoff of a moving window. Endpoints are handled by truncating the kernel and renormalizing the truncated weights to add to one.[1] The resulting low frequency trends capture the long-run fluctuations of the series. Second, we run a principal component analysis (PCA) on the estimated low frequency series and identify a common component. PCA is a statistical methodology that uses an orthogonal transformation to convert a set of possibly correlated variables (in our case time series) into a set of linearly uncorrelated variables called principal components. The idea is to identify one or more components that explain the largest-possible portion of the variance of the underlying series. If a single component is associated with an eigenvalue higher than one and explains a large fraction of the underlying variance, this component can be interpreted as a common factor driving variation in all series. In our case, we interpret the common factor as a measure of the long-run decline in labor market fluidity. With an estimate of the long-run trend in labor market fluidity in hand, we can then assess the magnitude of this decline and when it began.

1.1. The data series

The four quarterly series reflecting transitions into and out of employment (EU, UE, EN, and NE) are available from 1967:Q2 to 2015:Q3 for a total of 194 observations. All flows are expressed as a ratio of persons in the initial labor market state (e.g. EN shows the number of transitions from employment (E) to non-employment (N) as a share of E).[2]

[1] This approach has the advantage that it makes no assumption about reversion of the local mean. By contrast, the standard approach imposes mean reversion by using a stationary time series model to pad the series with forecasts and backcasts.

[2] Data since 2012 are available from the BLS gross flows statistics. Data through 2012 are from Elsby et al. [25]. Their data is derived from three sources. From June 1967 to December 1975, the data are made available by Hoyt Bleakley from tabulations put together by Joe Ritter. From January 1976 through January 1990, the data are made available from Shimer [61] on his website. From February 1990 and on, the data are available from the BLS gross flows statistics. Later in our

Regarding the annual series, we calculate aggregate job-to-job transitions from microdata for the Current Population Survey Annual Social and Economic Supplement (CPS-ASEC), as provided by the Unicon Research Corporation. Specifically, we calculate these transitions as the fraction of employed workers who report having had more than one employer in the previous year (respondents are explicitly instructed not to count multiple jobs held at the same time). The data are available from 1975 to 2014 for a total of 40 annual observations. Although this measure is a count of job transitions within a year, and therefore not conceptually identical to more common measures of month-to-month transitions, it is highly correlated with measures created by matching CPS cross-sections across months (e.g. Fallick and Fleischman [26]), which can be calculated from 1994 onward, as well as with job-to-job flows as measured in the Quarterly Workforce Indicators published by the Census Bureau, which are available from 2000 onward.[3] Interstate migration from 1975 to 2010 is from the IRS Migration Data. Because the methodology for measuring migration changed in 2011, we extend the IRS data post-2010 with growth rates of migration rates from the American Community Survey.[4] Finally, the job creation and job destruction data are from the Census Bureau Business Dynamics Statistics and are recorded from 1977 to 2013 for a total of 37 observations.

1.2. Results

Figure 1 shows the four quarterly series (EU, UE, EN, and NE) and the extracted low frequency components, while Figure 2 shows the series recorded at annual frequency (JtJ, IM, JD, and JC) and the

analysis, we estimate gross flows by demographic characteristics, which we calculate using monthly CPS data matched with code that Rob Shimer provides on his website.

[3]The correlation of our estimate with each of these other estimates is 0.97.

[4]For the years that the IRS and ACS data overlap, the level and changes in aggregate migration are quite similar (Molloy et al. [54]).

extracted trends. All four annual series show clear evidence of downward trends over the sample period. Similarly, UE declines for most of its (somewhat longer) sample period. EU increases from the mid-1960s to mid-1980s, but then falls for much of the remaining period. EN falls from the mid-1960s to the late 1990s, and then flattens out. Finally, of all these measures, NE shows the least evidence of a downward trend, although as described in the next section, this is because a rise in NE for prime-age females is offset by declines in NE for prime-age men and younger persons, consistent with trends in labor force participation for these groups. Broadly, the evidence emerging from Figure 1 and Figure 2 suggests a long-run decline in fluidity, with all trends at the end of the sample being below or well below their level in 1975.

Because the series in Figure 1 and Figure 2 have different scales, it is hard to compare the magnitude of the declines. Therefore, in Table 1, for the low-frequency component of each series we report the sample mean, the sample standard deviation, the minimum and the year in which it occurred, and the maximum and the year in which it occurred. In each case, the minimum is located at the very end of the series while the maximum is at the beginning of the sample period (with the exception of NE for which it is in the middle). On average, the size of the decline in fluidity measures is substantial. If we compare the deviation between maximum and minimum, the drop amounts to almost a fourth of the initial level for JtJ and to around 20 percent for IM, EN, and JC, while it is smaller for EU and NE. Also, these long-run fluctuations seem to be highly correlated. In Table 2 we report the pairwise correlations coefficients across the estimated eight trends. Although these correlations are computed over a relatively small number of observations (when using annual data), the evidence emerging from Table 2 suggests a high degree of comovment across the low frequency components of these labor market fluidity measures, with the exception of NE which appears less correlated with the other trends.

As a second step we investigate whether the aforementioned low frequency trends share a common factor by running a Principal Component Analysis (PCA). Table 3 reports results from three distinct PCAs: one using the trends based on the four quarterly flows, one using the trends in the five annual measures of worker flows, and one using the trends in all eight annual series. In all cases the first component explains the majority of the variance of the underlying series, so we focus on the first component only as our common component of interest. As the right-hand-side of Table 3 shows, the PCAs put a positive weight on nearly every variable in the analysis (again NE is the exception), indicating that the common component identifies a factor that is positively correlated with each of these flows. Finally, we plot the first components of the three PCAs in Figure 3. The three first components are very similar and they convey the same message: there is a clear downward long-run trend that is common across virtually all measures of labor market fluidity. Because the factors generated by PCA are normalized to have a mean equal to zero and variance equal to one, the resulting factor does not give much insight into the magnitude of the decline in fluidity. Taking a simple average of the eight individual long-run trends suggests that general labor market fluidity decreased by about 13 percent over the period that we examine.[5]

Finally, we address the timing of the decline in labor market fluidity. While various idiosyncratic factors may have caused each series to have peaks in different time periods, the PCAs shown in Figure 3 suggest that the common trend in labor market fluidity has been declining since at least the early 1980s. Notably, these declines appear to be fairly constant for most of the period that we consider. To a large extent, we obtain this result because we choose trends that filter out all but the very low-frequency

[5]A simple average is not a bad approximation since the PCA assigns roughly equal weights to most series. Using the factor loadings from the PCA as weights also yields a weighted average decline of about 13 percent.

movements in each series. Filters that allow for higher-frequency movements, such as 10 to 15 year cycles, are more difficult to interpret because they tend to be correlated with the severe business cycles of the early 1980s and the late 2000s. Thus, it is difficult to distinguish possible changes in the long-run trend from the fact that there were two severe business cycles towards the beginning and end of the period that we consider. Consequently, we focus on the lowest-possible frequency movements in labor market fluidity, smoothing through possible inflection points in the data.[6]

The above evidence is robust to using two alternative filters to estimate the low frequency movements of the individual measures of labor market fluidity: (i) the lowpass version of the Christiano and Fitzgerald [17] filter, and (ii) the low frequency cosines projection method suggested by Müller and Watson [56].[7] The results of these robustness checks are reported in Appendix A. Specifically, Table A.1 and Table A.2 mimic the results reported in Table 2 showing the correlation of the estimated trends using the Christiano and Fitzgerald [17] filter and Müller and Watson [56] method respectively. Also, Table A.3 and Table A.4 mimic the results reported in Table 3, showing the PCAs results using the estimated trends employing the Christiano and Fitzgerald [17] filter and Müller and Watson [56] method respectively. Furthermore, Figure A.1 compares the first component of our baseline PCA (the one run on low frequency components estimated using a biweight filter) with the ones run on the alternative set of trends. Overall, these robustness checks largely confirm the baseline evidence of a long run decline in fluidity which tends to be positively related to virtually all separate transition measures. As a final check, we also reversed the

[6]Decker et al. [22] and Decker et al. [23] emphasize that the decline in job creation and job destruction appears to have accelerated around 2000 in some industries. Although an analysis of the inflection points in specific industries is undoubtedly valuable in shedding light on the specific factors affecting labor market flows in these industries, we choose to focus on common trends that persist over the entire 30 to 40 years in hopes of shedding light on contributors to the decline in labor market fluidity that are broad-based across industries and pertain to a long period of time.

[7]The parameters of the Christiano and Fitzgerald [17] filter are set to retrieve cycles longer than 30 years. For the cosines projection method we use two cosines functions.

order of our two-step procedure (i.e. running the PCA first on the raw series and then estimating the low frequency trend of the common component). In this case we obtained similar results to the baseline, although without first smoothing out the cyclical fluctuations in these series, it is more difficult to identify a factor that has a positive weight on all measures of fluidity in the PCA.

2. Worker Demographics and the Decline in Mobility

Declines in labor market fluidity over the past three to four decades coincide with other demographic and economic trends that seem, on their face, like logical explanations for a substantial portion of the secular decline in fluidity. Examples include aging of the population and rising female labor force participation. Previous research has found these demographic shifts account for only a little of the secular decline in some measures of fluidity. For example, Kaplan and Schulhofer-Wohl [47] show that changes in the age distribution, changes in the types of occupations and industries, rising income inequality, and increased numbers of dual-earning households only explain a small amount of the decline in cross-state migration. Molloy et al. [55] show that, in addition to being unable to explain much of the decline in interstate migration, shifts in these and other demographic factors (e.g. education and geography) cannot explain much of declines in employment transitions across firms, occupations, or industries.

Similarly, Hyatt and Spletzer [45] show that changes in a variety of worker characteristics (e.g. age, gender, education) and firm characteristics (e.g. size and age) explain only a small fraction of the change in worker flows like job-to-job transitions, hiring rates, and separation rates. Regarding flows measured from the firm's perspective, Decker et al. [21] find that the shift in the age distribution of firms (towards older firms) can account for no more than a third of the decline in job creation and destruction since the late 1980s.

In this section we focus on fluidity as measured from the perspective of the worker in order to examine the amount of the aggregate change in fluidity that can be explained by changes in the distribution of demographic characteristics (e.g. aging of the population), as well as to identify important differences in fluidity across demographic groups. For ease of exposition, we examine on job finding rates (UE and NE transitions as a share of non-employment), job separation rates (EU and EN as a share of employment), and job-to-job changes (as a share of employment, although we also note where findings are different for the UE, NE, EU, and EN flows.

Demographic shifts should affect movements in labor market fluidity to the extent that average levels of fluidity vary across demographic groups. Indeed, Figure 4 reveals a number of important demographic differences in job finding and separation rates, as well as demographic differences in job-to-job transitions and migration (as measured in the March CPS). Job separation rates tend to be higher for younger workers (16-24 years). Job finding rates, and job-to-job transitions are higher than average for younger workers and lower than average for workers nearer retirement age (55+). Migration rates are also higher for younger persons, and lower for older persons. Since age appears to be an important distinction for many of these measures, the gradual aging of the labor force offers one potential, cohesive explanation for the decline in these measures of fluidity.

2.1. Changes in fluidity attributable to changes in demographics

To assess the contribution of shifts in the distribution of characteristics to the aggregate movement in these measures of fluidity, we estimate counterfactuals that hold the average fluidity over the period fixed for each demographic group and allow the population shares of each group to evolve as they actually did. Specifically, we estimate the following regression, which follows the approach in Moffitt [51]:

$$y_{ikt} = \beta_0 + X_{ikt}\beta_k + \Theta_t + \epsilon_{ikt}. \tag{3}$$

Here, k is an age-sex-education-marital status category. For ease of computation, we collapse the data to k-level cells by age, sex, education (no high school degree, high school degree, some college but less than a 4 year degree, 4 year degree or more), and marital status (ever married or not), by year.[8] Included covariates are, depending on the specification, a full set of age dummies, sex dummies, four education group dummies, and two marital status dummies. When we only include year fixed effects in the regression, the fixed effects estimate the average fluidity in each year. When other demographic controls are included, the year fixed effects represent the annual average fluidity in each year after controlling for the included demographic controls. We then normalize the year fixed effects to zero in the start of the sample (1976 for labor market flows, and 1981 for cross-state migration).

Figure 5 plots these fixed effects for job finding and job separation rates, job-to-job transitions, and interstate migration. The black solid line plots year fixed effects without controlling for any demographics; the trends in these series correspond with the aggregate trends shown in Figures 1 and 2. The orange line

[8] We estimate the regressions with weighted least squares, weighting by the size of each cell.

plots the year fixed effects after controlling for age, sex, and marital status. The blue line shows the year fixed effects after also controlling for education, another characteristic of the workforce that has displayed a considerable secular change over the past four decades.

For all series, the first set of demographic controls explain at most half of the decline in all measures of fluidity over this period. When we include education, we can *over-explain* the decline in job separations. Mechanically, persons without a high school degree tend to have high job separation rates, and persons with a 4-year degree or better have low job separation rates; these differences are so big that the rise in educational attainment would be expected to reduce job separation rates even more than actually occurred, all else equal. Meanwhile, shifts in the distribution of education do not do much to help explain any of the movements in job finding rates, job-to-job transitions, and migration.[9]

2.2. Changes in fluidity for particular demographic groups

As shown above, demographic shifts appear to explain some, but not all, of the secular decline in fluidity. To understand what may be responsible for the remainder, it is useful to consider differences in fluidity trends across demographic groups.

Returning to Figure 4, declines in the job finding rates have been much steeper for prime-age males and younger persons; job finding rates for prime-age females rose for much of this sample period. By contrast, separation rates fell more for older workers and prime-age females. The demographic differences in these

[9] As shown in Figure A.2 in Appendix A, aging can explain about half of the decline in EN, EU, and NE, but none of the variation in UE. Adding in education also helps explain more of the decline in EN, EU, and NE, and also provides no additional contribution to UE.

trends mirror well-documented differences in labor force participation (e.g. Aaronson et al. [1], CEA [16]). Female labor force participation rates have risen steadily through the late 1990s, reflecting changes in social and workplace norms and increased job opportunities for women. This pattern likely contributed to the secular increase in job finding rates and secular decline in job separation rates for prime-age women. Meanwhile, participation rates for prime-age men have been in a prolonged decline, likely in part due to technology- and globalization-driven shifts in labor demand away from male-dominated occupations and industries. The decline in participation is consistent with the secular decline in job finding rates and rise in job separation rates for this group. Participation rates for older persons have also risen as retirement ages have increased, and consequently their job separation rate has fallen. Finally, the decline in the job finding rate of younger persons is consistent the decline in participation rates for this group, likely reflecting increased rates of college-going.[10]

Trends in job-to-job transitions and interstate migration are much more similar across groups than job finding and separation rates (Figure 4). In particular, job-to-job transitions have declined for all groups except those 55 years and older, and cross-state migration rates have declined for all groups.

To summarize, there are important differences in labor market fluidity across groups of workers by sex and age. Shifts in the composition of the population toward groups that tend to make labor market transitions less often explain some, but no more than half, of the aggregate trends in fluidity. Thus, although demographic shifts clearly matter, there remains considerable room for other explanations. Some

[10]Demographic differences in the various component flows of job finding and job separation rates are consistent with explanations related to labor force participation (Figure A.3 in Appendix A). For younger workers and prime-age mails, NE flows have declined notably. For prime-age women, EN flows have fallen somewhat, and NE flows had rose through 2006 or so before dropping back during and after the recession. For 55+, EN flows have also fallen, likely reflecting later retirement ages. Also, and less likely to be related to labor participation decisions, EU flows have fallen a bit for most demographic groups.

of these explanations may relate to group-specific trends. In particular, some age-sex specific trends in job finding and separation rates appear to reflect secular trends in labor force attachment over this period, suggesting that some of the heterogeneity we observe in the overall decline in fluidity is related to a group's connection to the labor market. Hence, understanding broad changes in the labor market that affect labor force attachment is likely to be key to fully explaining the decline in fluidity. For job-to-job transitions and migration, the trend in fluidity is fairly similar for most groups, suggesting that explanations for this trend should apply to broad swaths of workers.

3. State-Level Differences in Labor Market Fluidity

As is true for any large country, the US labor market is a collection of smaller, local labor markets that differ along many dimensions. Geographic movement of workers and firms helps to integrate these markets, although this integration is far from perfect because long-distance migration is costly. The decline in the national average of labor market fluidity that we have documented thus far must therefore also occur at the local level, but perhaps to varying degrees across states. In this section, we analyze variation in trends in labor market fluidity across states in hopes of shedding light on the factors that are behind the decline in the national average. States are imperfect measures of local labor markets, as some states comprise several local markets whereas others are strongly connected to a labor market in a neighboring state. However, state-level trends in fluidity are likely similar to those at finer levels, making them a useful approximation of sub-national labor markets (Molloy et al. [54]). Moreover, data concerns are also relevant: publicly-available datasets with annual data on labor market transitions are too small to be able to identify geographic areas any smaller than states. For this analysis, we create state-level measures of fluidity using the same eight measures that we used in the aggregate analysis. Most measures are from the CPS: flows from employment to unemployment (EU), flows from employment to not-in-labor force (EN), flows from unemployment to employment (UE), flows from not-in-labor force to employment (NE), the fraction of workers who changed employers (JtJ), and the fraction of individuals that moved across state lines (IM). Data on job creation (JC) and job destruction (JD) are from the Business Dynamics Statistics produced by the Census Bureau. Due to the availability of the migration and job creation/destruction variables, the eight measures combined are available from 1980 to 2013.[11] We focus on annual rather than quarterly data because we are concerned that many states are too small to reliably estimate labor market flows at a higher frequency.

Our state-level analysis follows the same two-step procedure we use with the aggregate data. We start by estimating a state-level trend for each measure of labor market fluidity. With only 34 annual observations for each state, it is not possible to employ our time series techniques to estimate these trends. Instead, for each state we estimate a linear time trend from an OLS regression that includes a trend and the state unemployment rate (contemporaneous and one lag). The coefficient on the linear trend reflects the average decline in each measure by state, after (roughly) accounting for the business cycle.

In the second step, we combine the data for all states and use principal component analysis to identify a trend that is common across all of the eight measures of fluidity. The factor with the largest eigenvalue explains 70 percent of the variation among these eight trends and has a positive factor loading on each one. Thus, we use this common factor as the average decline in labor market fluidity by state.

[11] All results reported below are robust to measuring migration using the IRS data rather than the CPS; we prefer the CPS data for this purpose because the IRS data are not available for DC, Alaska and Hawaii.

One concern with this method is that the linear trend assumes that declines in fluidity have been constant over time. As discussed above, it is challenging to discern whether trends in labor market fluidity have changed during this time period because the recessions at the beginning and end of the period were more severe than the recessions in the middle. Thus, apparent changes in the trend might instead reflect a cyclical pattern. Nevertheless, to assess the possibility that the state-level trends have not been constant over time, we interact the linear trend with an indicator for the second half of the sample (post-1996). Because we estimate separate regressions for each state and each measure of labor market fluidity, this exercise yields 408 estimates (=51 states times 8 measures) of trend breaks. In only about one quarter of cases, the estimated trend changes by more than 20 percent from the first half to the second half of the sample and this difference is statistically significant at the 5 percent level or smaller.[12] Consequently, although there are clearly cases where the trend has not been constant over this 34 year sample period, we conclude that characterizing the general pattern with a linear trend is a reasonable approximation.

Figure 6 reports the trend in labor market fluidity for each state.[13] The first point to take away from the figure is that the estimated trend is negative in all states. Nevertheless, there is a substantial amount of variation in declines in fluidity across states. Declines are relatively mild in a number of eastern states like North Carolina and Connecticut, averaging around 0.5 percent of the initial level of fluidity per year,

[12]Trend breaks appear to be more common for job destruction, job-to-job transitions, and flows from employment to out of the labor force.

[13]These estimates are not dissimilar to those shown by Davis and Haltiwanger [19], who calculate trends for all 51 states as the difference in average job reallocation from 1988-1990 to 2008-2010. The correlation of our estimates with theirs is 0.64. Differences between the two sets of estimates owe to a number of methodological differences: we include a wider range of measures of labor market fluidity, our sample period is longer, we control for the business cycle, and we estimate trends using all annual data points rather than taking the difference of the end points.

whereas they average more than 1.5 percent per year in western states like New Mexico and Montana.[14]

The geographic pattern of declines in labor market fluidity is intriguing because states in the West differ from states in the East along a number of demographic and economic dimensions. With only 51 states and state-level characteristics that are highly correlated with one another, it is extremely difficult to tease out which state-level characteristics are robustly correlated with the decline in labor market fluidity. Nevertheless, as an attempt to examine this question, we estimate a series of regressions with the decline in labor market fluidity as the dependent variable and various sets of state-level characteristics as independent variables.[15] We consider the following sets of characteristics: population age, educational attainment, marital status homeownership, industry, occupation, union membership, and class of worker (private, self-employed or government).[16] For each set of variables, we consider correlations with the average level in 1977 to 1979 as well as the trend from 1980 to 2013. These trends are estimated using the same methodology as was used for estimating trends in labor market fluidity. The dependent and independent variables are all scaled to have a mean of 0 and standard deviation of 1, so that the magnitudes of the coefficients can be interpreted in terms of standard deviations.

For most sets of characteristics, at least one or two appear to be at least moderately correlated with the trend in labor market fluidity. One notable exception is homeownership—neither the initial level nor trend is correlated with the trend in fluidity—suggesting that changes in the cost of homeownership or

[14] In order to interpret this factor as an average percent change, we calculate the weighted average of the 8 trends in the individual fluidity measures and divide by the weighted average of the initial levels of these measures in 1980. In both cases, the weights used are those on the first factor of the PCA.

[15] We do not have enough observations to consider all potential state-level characteristics in a single regression. Although we could regress fluidity on each state characteristic individually, doing so would likely lead to a large number of spurious results because many state-level characteristics are mechanically correlated with other characteristics; for example, states with a large fraction of young people also have a small fraction of old people.

[16] All state characteristics except union membership are from the CPS-ASEC. Union membership was calculated from the monthly CPS and the Directory of National Unions and Employee Associations by Hirsch et al. [40].

the rise in homeownership are unlikely to explain the decline in fluidity. The remaining correlations are fairly difficult to distill into any clear explanations for the decline in fluidity, so we combine all of the variables that appeared to be meaningful into a single regression and drop variables that do not maintain a significant coefficient with a magnitude of at least 0.1 (i.e. a 1 standard deviation change in the variable is associated with at least a 0.1 standard deviation change in the trend in fluidity). Table 4 reports the results of this exercise. The coefficients for the full sets of variables are reported in Table A.5 in Appendix A.

Four interesting patterns emerge. First, declines in fluidity are smaller in states with larger initial shares of workers in administrative support and operator/fabricator occupations.[17] This correlation is likely related to the secular decline in demand for middle-skilled workers, which was particularly prevalent for workers in these categories (Autor [4]). We obtain similar results when only the trends in the shares of these occupations are included. However, when both the levels and trends are included, the estimated correlations with the trends become small and insignificant, likely because the correlations of the trends with the initial levels are over 0.9 for both variables.

The displacement of middle-skilled workers may generally contribute to additional churn in the labor market as these workers leave their old jobs and search for new jobs. It is worth noting that when we consider industry alone we also find that declines in fluidity were smaller in states with a higher manufacturing share, a sector where the change in demand for skill was more pronounced.[18] However, this result disappears once we control for the occupation shares. In general, these relationships suggest

[17]Occupations are defined using the 1990 categorization constructed by the IPUMS.

[18]Decker et al. [23] find that the trend decline in job reallocation (defined as job creation plus job destruction) is less steep in the manufacturing sector than in some other industries, like retail and services.

that displacement of middle-skill workers has partly offset the general decline in fluidity in states with concentrated employment in routine-intensive jobs. However, they do not explain why the general decline occurred. In fact, they suggest that the decline in fluidity would have been more severe in the absence of the change in the demand for certain types of skill.

A second interesting correlation is that declines in labor market fluidity are marginally smaller in states with a larger decline in union membership, which is consistent with the notion that the decline of unions has reduced the frictions associated with hiring and firing workers. A third point to draw from this table is that accounting for these state characteristics reduces the coefficients on the Census division indicators somewhat, but differences in the Mid-Atlantic, Mountain and West divisions are still material. Thus, the geographic patterns are not largely attributable to the wide array of observable state characteristics that we consider here.

Finally, neither the levels nor trends of the state's distribution of age or education are related to the subsequent decline in labor market fluidity. While we do find a positive correlation between the trend in the population age 35 to 44 and the trend in fluidity, this result is entirely driven by the fact that declines in labor market fluidity were largest in Alaska, and this state also experienced the largest decrease in population share of this age group. Thus, although these demographic factors were important in explaining the downward trends in some of the individual measures of fluidity, the result does not hold when combining all measures together. This difference makes sense because these demographics had opposite effects on different flows—for example the rise in labor force participation of older workers reduces job separation rates but contributes positively to job finding rates. Combining these flows and focusing on the common trends across all measures of fluidity reduces the roles of such demographic effects.

4. Why Is Fluidity Declining? Benign and Less Benign Explanations

As outlined above, the decline in labor market fluidity has been widespread and sustained, features which suggest that any explanation for the declines must also unfold over a long period and must apply to a broad range of workers and industries. The declines in fluidity may at first be unsettling, but until we know the cause, it is difficult to judge whether it is bad for workers or the U.S. economy generally. In this section, we consider two classes of explanations: some that are not likely to imply adverse consequences for workers or economic activity, which we call "benign", and some that are more likely to imply adverse consequences, which we call "less benign." By bringing together results from the literature and performing additional analysis, we assess several explanations in both categories. We later suggest several next steps based on the explanations that our preliminary analysis suggests might be good candidates for explaining the decline in fluidity.

4.1. Benign Explanations

Reduced transitions may reflect improvements in the worker-firm relationship, and thus less need for workers to change jobs. A major source of such improvement may be better matching between workers and firms. As matching improves, it becomes less likely that another job exists where a worker would be more productive and thus transitions in the labor market decline. A related, but separate cause of improvements in the worker-firm relationship could be if firms are investing more in their workers through increased training, thereby strengthening the ties that workers have to their firms. Finally, compensation may have become more responsive to changes in outside options for workers or to changes in productivity, reducing the need for the worker-firm match to dissolve in order for compensation to adjust to these forces. Although a full welfare analysis is outside the scope of this paper, all three of these explanations seem likely to be benign, if not beneficial, to the overall functioning of the economy. In this subsection, we explore the evidence for each of these hypotheses in turn.

4.1.1. Evidence on Improved Matching

If matching has improved and wages reflect match quality, then a worker's wages will be more closely aligned with his best possible match quality over the course of his career. We cannot test for the alignment between realized wages and potential match quality empirically since match quality is unobservable. However, if wages proxy for realized match quality, then trends in wages can provide some insight into the plausibility of the improved matching hypothesis.

To fix ideas, define match quality δ in the following way. Let Θ_F be the set of all firms and Θ_W be the set of all workers. For simplicity of notation, assume firms have only one worker. $M(\Theta_F, \Theta_W)$ is a one-to-one allocation of workers to firms.

Then a role for match quality implies the existence of an allocation M^* such that there is no Pareto-improving switch of workers across firms that would raise or hold constant match quality for all firms.

Specifically, under M^*, there is no change in worker firm matches k, j such that:

$$\delta^{fj}_{wk} > \delta^{fj}_{wj} \ \ and \ \ \delta^{fk}_{wj} \geqslant \delta^{fk}_{wk}. \tag{4}$$

We define improved match quality to mean that the labor market has moved closer to M^*. More firms employ their M^* worker, and more workers are employed at their M^* firm.

To consider the impact of improved matching on wages, assume wages equal match quality plus a base equal to the average level of human capital among workers (which we assume can be deployed for the same return in any firm):

$$\omega_w = \bar{h} + \delta^{f}_{w|M}. \tag{5}$$

When matching falls short of M^*, swapping workers can result in Pareto improvements in match quality that raise wages for some and without lowering wages for others. Improvements in matching should therefore result in higher average wages (all else constant) as these matches are more frequently made. To put it another way, under a better matched allocation, more workers are employed by firms at which, were they to change employment, the match quality for themselves, their replacement, or both would be lower.

The incidence of these higher wages over the course of a worker's career depends on when match quality is revealed in the worker-firm relationship. If match quality is revealed prior to starting employment and if match quality is rising, then we should observe starting wages rising over time. Hyatt and Spletzer [46] test for such a trend over the period 1996-2014 in the CPS and 1998-2008 in the LEHD. They find no evidence that starting wages have increased at all in these periods.

Because the decline in labor market fluidity predates the time period analyzed by Hyatt and Spletzer [46], we turn to data on three cohorts from the National Longitudinal Surveys to look at initial wages on a job. We assemble a panel of three cohorts the National Longitudinal Surveys (NLS): the NLS-Young Men (NLS-YM); the NLS-Youth 1979 (NLSY79), and the NLS-Youth 1997 (NLSY97). We focus on results for men because the labor force participation of women changed markedly over these three decades and we are concerned that female labor force participants in the late 2000s are different in many unobservable ways from their counterparts in the late 1970s, which complicates cross-cohort comparisons.[19] Because respondents in the latest waves of the NLSY97 are still young, we restrict each sample to respondents aged 22 to 33 to maintain comparability across the samples. Roughly speaking, our cohorts represent the labor market experiences of young workers during the early 1970s (the NLS-YM), the 1980s and early 1990s (NLSY79) and the 2000s (NLSY97).

To calculate starting wages, we regress the real wage of young, male workers who have less than 1 year of tenure at their current employer on indicators for age, race, education, and year. The regression is estimated separately for three cohorts over the following time periods: 1966 to 1981, 1979 to 1994, and 2002 to 2013. The constant of this regression reveals the average starting real wage in each time period. Table 5 shows that in this sample, starting wages rose somewhat from the first period to the second, but then decreased in the third period. If average starting wages constructed in this way reflect average match quality, this pattern suggests that matching was better in the 1980s and early 1990s than in more recent

[19]This sample is similar to one constructed for Molloy et al. [55], but we have made it publicly accessible by omitting use of restricted geocoded variables. We have also updated the data construction in a number of other ways. Details of the data assembly available upon request from the authors.

years, a result that is inconsistent with a rise in match quality.[20] The same pattern holds within broad skill groups, suggesting no trend improvement in initial match quality even among workers who faced rising demand for their skills over this period. The data seem more likely to indicate that starting wages fluctuate modestly with no real trend over time. We find similar results in the PSID, for which we can look at older workers as well as younger workers, but only from 1976 onward (see Figure A.4 in Appendix A).

If match quality is only revealed after a worker has been with a firm for some amount of time (i.e. it is an experience good), the quality of *retained* matches should rise across cohorts even though the quality of new matches would not improve. As long as wages reflect match quality, returns to tenure with an employer should rise across cohorts of workers. We also test this hypothesis empirically using our panel of young workers from several cohorts of National Longitudinal Survey respondents. To examine changes in returns to employer tenure across cohorts, we estimate the following wage equation:

$$\begin{aligned} y_{it} &= \beta_0 + \beta_1 indten_{it} + \beta_2 indten_{it}^2 + \beta_3 occten_{it} + \beta_4 occten_{it}^2 + \\ &\quad + \beta_5 jobten_{it} + \beta_6 jobten_{it}^2 + \theta X_{it} + \Theta_t + \Theta_i + \epsilon_{it}. \end{aligned} \tag{6}$$

The dependent variable is log real hourly wages for respondent i on the main job in survey year t. The hourly wage is the "hourly rate of pay" variable constructed for each reported job by NLS administra-

[20]The estimation controls for year effects, but not across rounds of the NLS, so one may still worry about cyclical differences across the three NLS cohorts. However, the average unemployment rate was 4.3 percent in the first period, 7.0 percent in the second period, and 6.5 percent in the third period. Therefore, it is unlikely that changes in the cyclical position of the economy over these three cohorts are obscuring a secular increase in starting wages.

tors. X_{it} is a set of additional background controls that includes age, age squared, and four educational attainment dummies (dropout, high school graduate, 1-3 years of college, 4+ years of college). Θ_t is a set of survey year dummies. Θ_i is a set of person fixed effects, which are included to mitigate the concern that higher-quality workers may stay longer with an employer, biasing up the estimated return to tenure. It is worth emphasizing that the regression includes controls for occupation and industry tenure, so the estimated return to employer tenure does not include returns to more general human capital that a worker can take with them when they change employers.

We estimate this equation separately for each of our NLS cohorts. The results are reported in Table 6, which shows estimates of the returns to a third year of tenure with an employer. We focus on the third year of experience because average tenure in each sample is around two to three years. Returns to employer tenure are economically small and insignificant for all three cohorts, providing no evidence for rising returns to employer tenure over time.[21]

4.1.2. Changes in Employer-Provided Training

Cairo [10] and Cairo and Cajner [12] develop models in which job-related training reduces the propensity of workers to separate from employers. Thus, an increase in training requirements for broad groups of workers could contribute to the secular decline in labor market fluidity. Based on a model simulation, Cairo [10] concludes that rising training requirements can account for about one third of the multi-decade

[21] Our estimated returns to tenure are smaller than many others in the literature (such as Topel [65] and Buchinsky et al. [9]) because we are controlling for occupation and industry tenure (Parent [57]). Our estimates are similar in magnitude to those reported by Altonji et al. [3], who model wages, employment transitions, and hours jointly for men in the PSID from 1975 to 1996.

decline in measured job turnover in equilibrium. A related hypothesis is that skill has begun to decay at a faster rate, as might be the case with firm- or job-specific skills. Fujita [27] proposes a model in which there is a secular increase in the risk of experience depreciation during an unemployment spell for all workers in an economy. Workers therefore become increasingly reluctant to separate from their firms and risk the loss of skill that would result from a failed transition to a new job. He argues that such a model can reconcile declining labor market turnover with stagnant wages and rising public anxiety about job security.

A challenge for the view that the decline in fluidity can be explained by an increase in job-related training is that evidence for a sustained increase in such training is limited. Cairo [10] finds that the share of workers employed in training-intensive occupations, as classified by the Dictionary of Occupational Titles (DOT), has increased over time. Periodic updates to the DOT's classification system also reflect more training within occupations in more recent waves (Cairo [10]). On the other hand, several studies that use direct evidence on the provision of training by employers finds no evidence of a sustained upward trend. In fact, it appears to have declined after peaking in the mid-1990s, a period with a number of Federal incentives for employer-provided training. This evidence is reviewed in Cappelli [14] and House [41]. Cairo and Cajner [12] also find that the incidence of formal on-the-job training was fairly flat, on net, from 1987 to 2007. Finally, it seems likely to us that a greater incidence of on-the-job training should result in greater returns to employer-specific tenure, for which we found no evidence in our analysis above.

4.1.3. Enhanced Compensation Flexibility

A third benign possibility that we consider is that reduced transitions reflect fewer frictions to wage or compensation adjustment. If firms are able to adjust compensation to reflect changes in productivity, this could reduce the need for layoffs as well as decrease voluntary separations in which workers leave a job in search of a larger wage adjustment.[22]

There is some evidence for this kind of enhanced compensation flexibility in the literature. In a set of papers on this question, Gottschalk and Moffitt [32] and Gottschalk and Moffitt [34] demonstrate that the transitory component of men's earnings rose between the late 1970s and the late 1980s. They argue that potential explanations for increased transitory fluctuations may be that enhanced compensation flexibility, perhaps arising from changes in worker protections or regulation, or from a more competitive product market that led to more rapid shifts in wages. Comin et al. [18] use COMPUSTAT data to test for an increasing correlation between firm level volatility in total sales and firm average compensation over time. They find that firms with higher sales volatility also exhibit higher average wage volatility, and that this relationship becomes much stronger after 1980. They argue that this change over time reflects enhanced pass-through of productivity fluctuations to worker wages. They further show that the compensation-sales volatility relationship is strongest in large firms. With a secular shift of employment towards larger firms, as documented by Davis et al. [20], one would expect the average pass-through of productivity to wage volatility to have increased.

[22]The implications of more frequent wage adjustment may not be viewed by workers as entirely benign if they dislike compensation volatility. The negative effects of greater compensation flexibility seem unlikely to outweigh the benefits of preserving a good match and reducing turnover costs, but a full welfare accounting of this channel is beyond the scope of this paper.

However, other evidence raises questions about the role of compensation flexibility in declining fluidity. First, there is little evidence of a sustained rise in the variance of transitory earnings. More recent analyses show a large increase around the late 1970s/early 1980s, followed by a long stable period and possibly even a reversal before it rose again into the 2000s and particularly in the Great Recession (Gottschalk and Moffitt [33]; Shin and Solon [62]; Koo [50]). One recent view holds that the increase in earnings volatility among men is related to severe recessions (Koo [50]), and is driven in large part by spells of unemployment (Ziliak et al. [66]; Koo [50]). Additionally, any increase in earnings volatility is confined to male workers, since earnings volatility has trended down for women since 1970 (Dynan et al. [24]).

4.2. Less Benign Explanations

Another class of explanations associates declines in fluidity with an increase in some cost of making an employment transition. There are a number of channels through which this might have happened—some might originate outside the labor market, such as through the housing market as job changes frequently require geographic transtions, whereas others might originate inside the labor market, for example through changes in firm or worker search behavior. In this section, we review a number of factors that may have led to an increase in the cost of changing jobs. In general, we think that rising costs are unlikely to be benign in their overall impact on the economy. Not only does a cost require resources to surmount, but a rise in transitions costs and the consequent reduction in reallocation will result in a less optimal allocation of resources.

4.2.1. The Role of an Aging Workforce Revisited

We begin by exploring the possibility that an aging workforce has led to fewer transitions in the labor market. Above, we showed that changes in the age composition of the population can explain a portion of the declines in some labor market flows—particularly those related to labor force participation—but the age distribution alone could account for less than half of the general decline in fluidity. However, simple decompositions might not yield the entire effect of the age distribution because if an aging workforce has broader general equilibrium effects on fluidity in the labor market, then aging could cause declines in fluidity even for older workers. For example, Shimer [60] develops a model in which a larger fraction of young workers generates more churning in the labor market, and older workers benefit from this churning as well. Similarly, Karahan and Rhee [48] develop a model in which an increase in the fraction of workers with higher moving costs (i.e. older workers) causes firms to hire more local workers, reducing the migration rates of all types.

To evaluate the likelihood of such general equilibrium effects, we look to see whether states with a larger decline in the fraction of young people have also experienced a larger decline in the labor market fluidity of older workers. Although we found little evidence of this correlation in Section 3 after controlling for other state characteristics, that analysis did not directly address the correlation of the youth share with fluidity of older workers. Consequently, we slightly alter the method described in Section 3 to measure declines in fluidity for older workers: we calculate state-level fluidity measures only for 35 to 64 years olds, excluding job creation and destruction since those two variables are not available by age of worker. Next we calculate the trends in these six measures using the same regression method and then combine the six trends using principal component analysis. Results of the principal component analysis are similar in that all variables have a positive loading and the first factor explains a large fraction of the variation in the data.

Figure 7 graphs the estimated declines in youth share against the estimated decline in labor market fluidity among older workers. The correlation is very weak. Correlations are similarly-weak for each separate measure of labor market fluidity, as well as for the simple average of the trend of each of these six measures (available upon request). Results are also similar when we control for the state characteristics that were found to matter in Section 3.

This evidence casts doubt on the idea that the decline in population of young workers has had a general equilibrium effect on the labor market transition rate of older workers. These results may be somewhat surprising in light of the analysis presented by Davis and Haltiwanger [19], which shows that changes in the fraction of young people in a state can predict changes in labor market fluidity in that state. Our analysis differs from theirs in at least two important respects. First, their measure of fluidity is not specific to an older age group, so some of the relationship that they find might be due to the direct compositional effect that younger workers tend to change jobs more often. Second, their correlations are estimated from 3-year changes in the dependent and independent variables. Thus, they may largely be picking up correlations at the business-cycle frequency rather than correlations among longer-run trends.

4.2.2. Declining Social Capital

Social institutions, and social capital in particular, are positively related to economic performance

(Knack and Keefer [49]).[23] Recent work argues that this relationship is causal, with greater aggregate social capital leading to improved long-run growth at the country level (Algan and Cahuc [2]). It is also possible that social capital is important for job and worker search, as there is evidence that jobs are often found are often made through personal networks (Bayer et al. [6]; Hellerstein et al. [38]; Hellerstein et al. [39]). Two major social capital measures for the U.S., both taken from the General Social Survey, have been declining for the last several decades (Glaeser et al. [30]). Declines in social capital—particularly the extent and strength of social networks—may raise the cost of job search by forcing workers to rely on more formal channels with less detailed information on the types of jobs available and the associated firm environments. In addition, or alternatively, reduced social capital may increase the cost of new hires as managers have to choose from workers about whom they have less personal information.

We use the restricted General Social Survey (GSS) with state identifiers to test for a relationship between social capital and fluidity in our state-level framework. The GSS has been widely used to measure social capital in the U.S. Of several such measures that can be constructed, the indicator variable for agreement with the statement "Most people can be trusted" is available over the longest period, for almost all years from 1972 to 2014. As shown in Figure 8, the fraction of people who agreed that most people can be trusted has declined markedly over the past three decades. According to Glaeser et al. [30], this is a useful measure of aggregate social capital as it indicates whether a community has a large share of members who are likely to behave in a trusting manner in their transactions. A second common measure of social capital from the GSS is the number of different types of membership organizations to which a respondent belongs. Glaeser et al. [30] prefer this as a measure of individual level social capital. We

[23] Social capital refers to the density of positive interpersonal relationships (connections) between members of a group.

focus on the trust measure because the memberships variable is not reported after 2004, and is not asked consistently in the years before that, but the results below are broadly similar across the two measures.

Figure 9 shows the relationship between a state's trend in fluidity and its trend in social capital as measured by the trust share.[24] Due to gaps in state coverage in the GSS from year to year, we can only reliably estimate trends for 41 states. Nevertheless, the figure shows a roughly positive relationship between changes in a state's social capital and its change in fluidity. States with large declines in social capital also saw larger declines in fluidity. A regression using these 41 points shows that this relationship is not statistically significant and is small in magnitude. In particular, a one standard deviation more negative trend in trust is associated with only a 0.06 standard deviation larger decline in labor market fluidity. This correlation more than doubles, to 0.15, when two outliers where trust increased substantially are excluded. It is also worth noting that some of the states with the largest declines in trust were in the Western Census region, and in this part of the country there were also unusually large declines in fluidity that were not explainable by standard worker characteristics or broadly defined industries. Moreover, in unreported results, adding the trend in trust to the state panel regression of Table 4, the resulting coefficients on the trend in trust are little changed from the univariate regression. Thus, this evidence is weakly suggestive that institutional changes, particularly a decline in social trust, may accompany the decline in fluidity. Whether this reflects the work of a third factor on both trust and fluidity, or whether it reflects a causal relationship, is impossible to tell but deserves further consideration by researchers.

[24]The trend in fluidity is the same as the one created in Section 3. The trend in trust is constructed using the same method as the trend in fluidity: the coefficient on a linear time trend in a regression controlling for the state's unemployment rate.

4.2.3. The Formalization of Hiring

A third possibility that we consider is whether the formalization of hiring processes has created additional rigidities in the labor market, reducing fluidity. To this end, we have assembled data on membership in the Society for Human Resource Management (SHRM), the major professional organization for human resource workers in the US. As shown in Figure 10, the fraction of the labor force who were members of the SHRM has risen substantially over the past 50 years, with the sharpest increases occurring from 1998 to 2008.

As above, we look to see whether states with larger increases in SHRM membership experienced larger declines in labor market fluidity. We estimate state-level trends in SHRM membership using the same regression framework described above. The SHRM data are only available by state for 1998 onward, so we estimate trends in labor market fluidity and SHRM membership only for this period.[25] As shown in Figure 11, there is little correlation between these two trends.[26] Therefore, it seems unlikely that the formalization of hiring has been the leading cause of the decline in fluidity over the past three to four decades.

4.2.4. Regulation of Land Use and Business Practices

A fourth candidate explanation we consider is regulations on businesses and construction. Specifically, we examine whether the regulation of land use, which restricts housing supply, or various federal regulations that could affect the costs of hiring or firing workers are associated with declining fluidity.

[25] The correlation between state-level trends in fluidity estimated from 1998 onward and estimated over the whole sample (1980 onward) is 0.64.

[26] This graph omits DC, which is an extreme positive outlier.

Restrictive land use regulations may be preventing the geographic reallocation of workers, and thus reducing labor market fluidity more generally (Ganong and Shoag [29]). Although this hypothesis may seem unlikely given that labor market fluidity has also declined substantially for transitions that do not require a change in residential location (Molloy et al. [55]), it is possible that it could be relevant if geographic reallocation is important for overall economic growth, as argued by Hsieh and Moretti [43].

Figure 12 displays the correlation of state-level declines in labor market fluidity with the average degree of regulation as measured by the Wharton Residential Land Use Regulatory Index (Gyourko et al. [36]), which is based on a survey conducted in the early 2000s. Panel data on regulations only exist for a handful of locations in the US, so it is not possible to estimate trends in regulation by state. The figure shows no support for the hypothesis that declines in labor market fluidity are more concentrated in states with tighter land use regulation.

Regarding industry-level regulation, Goldschlag and Tabarrok [31] show that job creation and job destruction are not, in fact, lower in industries with a higher degree of federal regulation in a panel of industries (3-digit NAICS) from 1999 to 2011. Their measure of federal regulation covers a broad range of regulations, some pertain directly to labor regulations, whereas others pertain to environmental regulation, national security, food safety, or other issues. They also show that federal regulation has been rising faster for manufacturing than for other broad industry categories since 1975, whereas fluidity has been declining by less in this sector (Decker et al. [23]). Moreover, among manufacturing industries, those with a larger increase in regulation did not experience a different trend in startups than those with no increase in regulation. Thus, the evidence in Goldschlag and Tabarrok [31] does not support the notion that declines in fluidity are driven by federal regulations. A related question is whether occupational licensing requirements, which have become considerably more common since the 1950s, has contributed to declining fluidity. Molloy and Wozniak [53] find no evidence that licensing is related to geographic mobility at the state level, which casts doubt on licensing as an explanation for declining fluidity more generally.

4.3. Implications of Declining Fluidity

Regardless of the cause, less fluidity in the labor market leads to fewer opportunities for workers to renegotiate their current employment arrangements using outside options as leverage or to change jobs.[27] In a key paper, Beaudry and DiNardo [7] argue that we can observe the results of such renegotiations by testing for the impact of labor market conditions on wages over the course of a worker's employment with a firm. To paraphrase their central claim: if broader market conditions at a given point in time affect a worker's wages, then the worker must have had an outside option she could credibly threaten to accept at that time. This gives worker the power to bargain for wage increases whenever outside offers exist and the worker could realistically accept them. Thus, they argue that in a spot market for labor, wages should be related to contemporaneous labor market conditions. On the other hand, if wages are determined by long-term implicit contracts between workers and firms, then contemporaneous conditions should have little effect. Rather, the relationship between wages and labor market conditions should depend on worker opportunity or ability to move across firms (mobility). If workers have limited mobility across firms, then wages are set at the start of a new worker-firm relationship, and wages should reflect labor market conditions at the time the worker was hired. By contrast, if workers have perfect mobility

[27]This holds even under the benign scenarios, as better matching or enhanced compensation adjustments make it less likely that workers obtain a credible outside option with which to bargain.

across firms, then the contract is reset whenever workers receive a better outside option, in which case wages should be related to the best labor market conditions since the worker was hired.

Using data from the PSID and CPS in the late 1970s and early 1980s, Beaudry and DiNardo [7] find the strongest support for the implicit contract model with perfect worker mobility. Grant [35] finds similar results using the original cohorts of the National Longitudinal Surveys and the NLSY79. We build on these studies, by examining how these relationships have changed over the past three decades. In particular, we estimate a log wage equation that includes labor market conditions at three points in time: contemporaneous conditions, conditions at the time a worker started her current job, and the most favorable conditions that obtained from the time the job started to the present. We use the annual national unemployment rate for all individuals aged 16 and older as our measure of labor market conditions.[28] Other controls include age, age squared, employer tenure, and employer tenure squared. We estimate the model in the PSID, the NLSY and the CPS (surveys that include the tenure supplement). The PSID and NLSY specifications also include individual fixed effects, while the CPS specification includes educational attainment, indicators for non-white and ever-married, and industry and region fixed effects. The one notable difference between our specification and that in Beaudry and DiNardo [7] is that our samples are long enough to include a quadratic time trend, so that our results are not driven by trends in unemployment and wages.

We find evidence that the role of external labor market conditions in wage setting has changed, at least since the 1990s. As shown in Table 7, like the earlier two papers, we find that the minimum unemployment

[28] Results are similar if we use state-level labor market conditions, allowing us to include year fixed-effects in the regression. However, we prefer the specification that uses national conditions because wage offers can come from outside of one's state of residence.

rate since a worker was hired had a large impact on wages in the 1980s and into the 1990s. This is true in all three datasets, regardless of the age range included. However, in the 2000s the connection between wages and the minimum unemployment rate is much weaker.[29] At the same time, initial conditions seem to have become more important for wages, although this correlation is not significant in our smaller data samples (the NLS97 or for young workers in the PSID). Thus, it seems that in the 2000s workers and employers renegotiated wage contracts less frequently with improving market conditions than they did in the 1980s and 1990s, a result that is consistent with the secular decline in labor market fluidity.[30] The question of when and by how much worker compensation adjustments happen is a key area for future research that we take up below.

Declining fluidity may have other impacts as well. For example, Davis and Haltiwanger [19] find that declining job creation and job destruction rates are linked to declining employment rates. Falling employment could lead to higher rates of involuntary unemployment. Alternatively, declining fluidity could make workers reluctant to separate voluntarily, leading to "precautionary" job holding and again an increase in the likelihood that the unemployed are there involuntarily. We investigate the possible link between fluidity and involuntary unemployment using the same state panel trend regression approach from previous sections. We regressed the cycle-adjusted state trend in involuntary unemployment on the state trend in fluidity.[31] We find that states with larger declines in fluidity saw higher shares of their unemployed who

[29]In the PSID, when the 2007-2009 recession and post-recession years are excluded, the coefficient on the minimum unemployment rate in the 2000s falls to -0.017 and is insignificantly different from zero. Otherwise, all estimates in Table 7 are robust to excluding that recession, as well as to omitting individuals whose current job has lasted less than one year (for whom initial conditions, best conditions and contemporaneous conditions are all the same).

[30]If compensation has become more flexible, then wages with the current employer may adjust more frequently than at the business cycle level. This could explain the decreasing importance of UR[min] for wages, but it is difficult to reconcile with a greater role for UR[initial].

[31]Involuntary unemployment was measured using the "whyunemp" variable in the CPS monthly data available on IPUMS.org for 1976 to 2015.

said an involuntary separation was behind their unemployment. The relationship was substantial: a one standard deviation decrease in fluidity was associated with a one-third of a standard deviation increase in involuntary shares of unemployment. The relationship was significant, with a p-value of 0.06.

5. Concluding Discussion: What have we learned, and what should future research tackle next?

Is the US labor market becoming less fluid? Previous studies have identified separate declines in various measures of labor market transitions over the past several decades. The accumulation of this evidence has led economists to ask if these separate findings represent a more general shift toward fewer transitions within the US labor market. Motivated by this question, we first seek to demonstrate a statistical connection between various measures of labor market transitions. To this end, we construct a unique measure of the trend in labor market fluidity by combining trends on the major flows into and out of employment with job-to-job transitions, interstate migration, and job creation and destruction. We then use principal component analysis to show that a single factor drives a majority of the long-run trends in each series. The identification of this common factor is the first contribution of this paper, as our approach demonstrates a statistical connection between the declines that have for the most part only been observed and analyzed separately in previous work. Our analysis suggests that labor market fluidity has declined 10 to 15 percent over the past $3\frac{1}{2}$ decades, indicating that this trend has been sizeable.

One advantage of our measure of fluidity is that it extends over a long time period, from the late 1960s to the present in its longest version, which allows us to investigate when the decline in fluidity began.

The data suggest that the declines in labor market fluidity began at least in the early 1980s, and perhaps in the 1970s. The result that this trend has persisted for at least three decades suggests that the causes of this trend also must have persisted for a long period of time. Dating the start of the decline is key to ultimately understanding its causes.

We devote the remainder of the paper to trying to understand these causes. Although we are ultimately unable to identify a clear reason for the decline, we make progress along several key dimensions. First, we first show that demographic changes can only explain a limited portion of the general decline in fluidity. Changes in labor force participation and educational attainment are relevant for some types of transitions and some demographic groups, but overall the general patterns are similar for most types of workers that we examine.

Next, using state-level variation in trends in labor market fluidity we find that fluidity is unrelated to most worker characteristics in the state as well as to the industrial composition of the state. One interesting exception is that states in the Mountain and Pacific Census Divisions have experienced larger declines in fluidity, even conditional on a wide variety of state characteristics. We also find that initial shares of middle-skill jobs in a state are negatively correlated with the decline in fluidity. It seems possible that the displacement of routine-intensive jobs may have increased labor market transitions for these workers, dampening the general decline in fluidity.

Finally, we consider a number of concrete explanations for declining fluidity, grouped into explanations with benign implications for the aggregate economy and explanations with less benign implications. The benign explanations that we consider are improved matches between workers and firms; enhanced flexibility in compensation that ties compensation more directly to productivity; and more intensive employer-provided training. The less benign explanations that we consider are sclerosis as a shrinking fraction of young workers reduces the liquidity of the labor market for workers of all ages, declines in social capital that make hiring and job search more difficult, the formalization of hiring practices, and an increase in regulatory barriers to labor transitions. Although our approach to assessing these explanations is descriptive and, in some cases, relies on the previous literature, we conclude that most of these potential channels are unlikely to explain the decline in fluidity. One exception is that states with a larger decrease in the fraction of people who report that strangers can be trusted tend to have experienced larger declines in labor market fluidity, suggesting that explanations related to social capital and networks are worth exploring in future work. We also believe the question of whether compensation adjustment within and across jobs has changed deserves more attention, and we return to this point below.

Although the evidence on potential explanations in this paper is far from definitive, in general we find little role for explanations that are related solely to worker characteristics or to general labor market institutions. Consequently, research into the connection between firm characteristics and declining fluidity seems like a promising avenue for future research. Although we can rule out the simple effect of industrial composition (other studies have found a limited role for industrial composition to explain the decline in job creation and destruction, including Hyatt and Spletzer [45] and Decker et al. [23]), there are many other firm characteristics that we are unable to explore with our data. For example, Decker et al. [21] and Davis and Haltiwanger [19] show that the secular decline in job creation and destruction is at least partly related to a decline in the number of smaller and younger firms. The research on the role that firm size and age may play in declining fluidity measures has so far focused on compositional effects, which may not account for all the ways in which these characteristics affect job turnover. The decline in new firm formation dates from the 1970s (Pugsley and Sahin [58]), which aligns well with the timing of the downward trend in fluidity. Therefore, we argue that more detailed examination of changes in the way firms and workers interact – particularly across firm size and age groups – would be quite valuable.

Getting inside the changing black box of the employment relationship also seems likely to be central for understanding declining fluidity. A series of recent papers document the important role of firms in rising earnings inequality (Card et al. [15]; Barth et al. [5]; and Song et al. [63]). Namely, a substantial portion of the widening in earnings inequality over the last several decades is due to a growing dispersion of earnings across firms, rather than increases in dispersion within firms. Could this rise in firm effect dispersion contribute to declining fluidity? This seems possible if rising disparity in pay by firm extends search time for workers, who may take more time between transitions as they look for a small number of high-paying employers. However, the evidence assembled so far suggests that the rise in earnings inequality is unlikely to be behind the decline in fluidity. First, the secular rise in earnings inequality has been linked to the decline in demand for middle-skilled workers, and if anything it seems like the changes in demand for skill have damped the long-run decline in labor market fluidity. Addtionally, we found no role for state-level earnings inequality when they were added to our state panel regressions in Section 3.

To examine this relationship further, we requested special tabulations of establishment level wage inequality from the National Compensation Statistics office of the Bureau of Labor Statistics.[32] As summarized in Table 8, the statistics that they provided also show little support for a connection between inequality across firms and labor market fluidity. In particular, the rise in the 90-10 differential in average compensation across establishments (a proxy for firms) is largely a coastal story. Middle, non-coastal

[32] We thanks Brooks Pierce and Jesus Ranon of the National Compensation Statistics program, and the Bureau of Labor Statistics, for tabulating these data for us.

Census divisions saw little increase in establishment-level pay inequality. This pattern does not align well with geographic pattern that we observe in labor market fluidity. Thus, an increase in firm heterogeneity seems unlikely to explain declining fluidity, although it is possible more could be done here.

A more promising avenue for research would be an exploration of the wage changes that workers experience both within and across firms. For example, it would allow a cleaner assessment of whether within-firm earnings volatility has increased, possibly signalling the stronger connection between compensation and productivity discussed above. It would also allow for an examination of whether the return to changing employers has fallen, which might have occurred if large firms offer a less variable set of contracts to a given worker.[33] Enhanced information about firms would also allow an exploration of how firm output volatility relates to hiring and separation, and whether these relationships have changed over time. Simply documenting changes in compensation within and across firms in a robust way over time would be a helpful step because there has been very little work on this area to date.[34]

However, the data challenges in most publicly-available datasets are substantial. Therefore, use of matched employer-employee data, which tend to be large and high quality, would be an appropriate resource to explore. Although it is also often the case that demographic information is more limited in these datasets, our analysis suggests that demographics can largely be set aside, so omitting them from this type of analysis would likely not be a severe impediment. On the other hand, many such datasets are only available for recent decades, so one would need to extrapolate from this evidence to the entire three

[33] Cannon et al. [13] find that compensation differences across firms but within occupation groups declined from the early 1980s to the late 1990s. However, they only examine average compensation differences, not changes in the compensation received by workers that change employers.

[34] Molloy et al. [55] attempt to document changes in the returns to employer changes over time using three cohorts of the National Longitudinal Survey (NLS). An updated version of that analysis, which includes data through 2013 and the age groups in this paper, shows that returns to employer change in fact have changed little for cohorts of younger workers since the 1970s. Returns to an employer change were about 3-4 percent in each NLS cohort.

to four decades over which labor market fluidity has been declining. Finally, the analysis in this paper highlights how little we know about secular changes in the terms of employment, such as information on screening and hiring practices, as well as firm-provided training. More information here would certainly be useful going forward. And enhanced matching across US government data sets might help us understand more about how firms use firing versus other separations.

Finally, another important topic for future research is a clearer understanding of the effects of the secular decline in labor market fluidity. We have shown that this trend appears to have coincided with a reduced frequency of wage renegotiations between employers and workers. The reduced responsiveness of wages to external labor market conditions may signal additional rigidities in the compensation setting process, or it might signal a diminished need to renegotiate. Davis and Haltiwanger [19] show that decreases in worker and job reallocation are associated with lower employment rates of the young and less educated. We take a further, but still preliminary, step and examine the relationship with declining fluidity and involuntary unemployment, and find a strong negative relationship. This early evidence strongly suggest that the first-order effect of the aggregate decline in fluidity is unlikely to be benign. More investigation into whether other, non-benign outcomes are linked with the decline in fluidity would be a useful next step toward understanding these effects.

Although this paper has raised at least as many questions as it has answered, we hope that it has made a few things clear. Labor market fluidity has been declining since at least the 1980s, and has been fairly broad-based across types of workers broad industrial sectors. Geographic differences in the extent of the decline in fluidity suggest that the causal factor may not be equally present across the US economy, but investigations into causes that are readily explored with publicly available data turned up no leading explanation. Because this trend has persisted for so long and touches on so many types of workers and firms, more research on the causes and consequences of this trend would be extremely valuable. We advocate further research using administrative data, in particular firm-level data that can inform our understanding of these changes.

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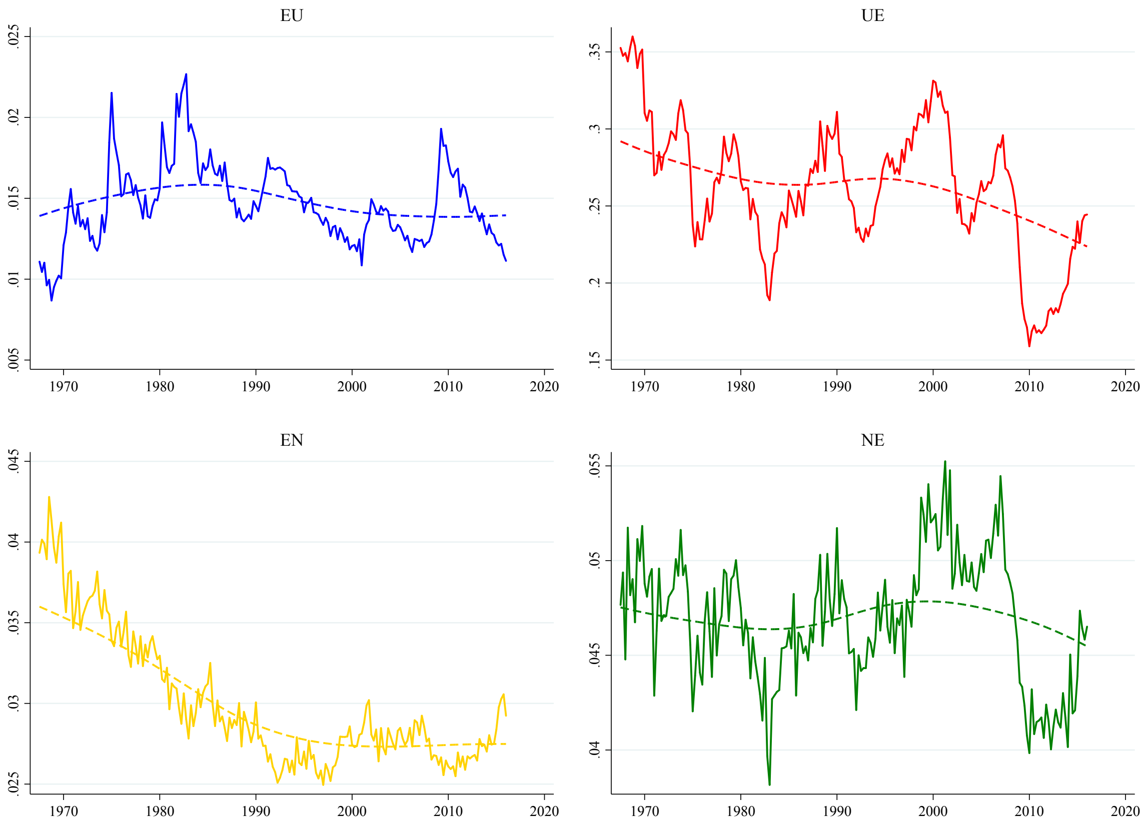
Figures and Tables

Figure 1: Quarterly series and estimated low frequency trends.

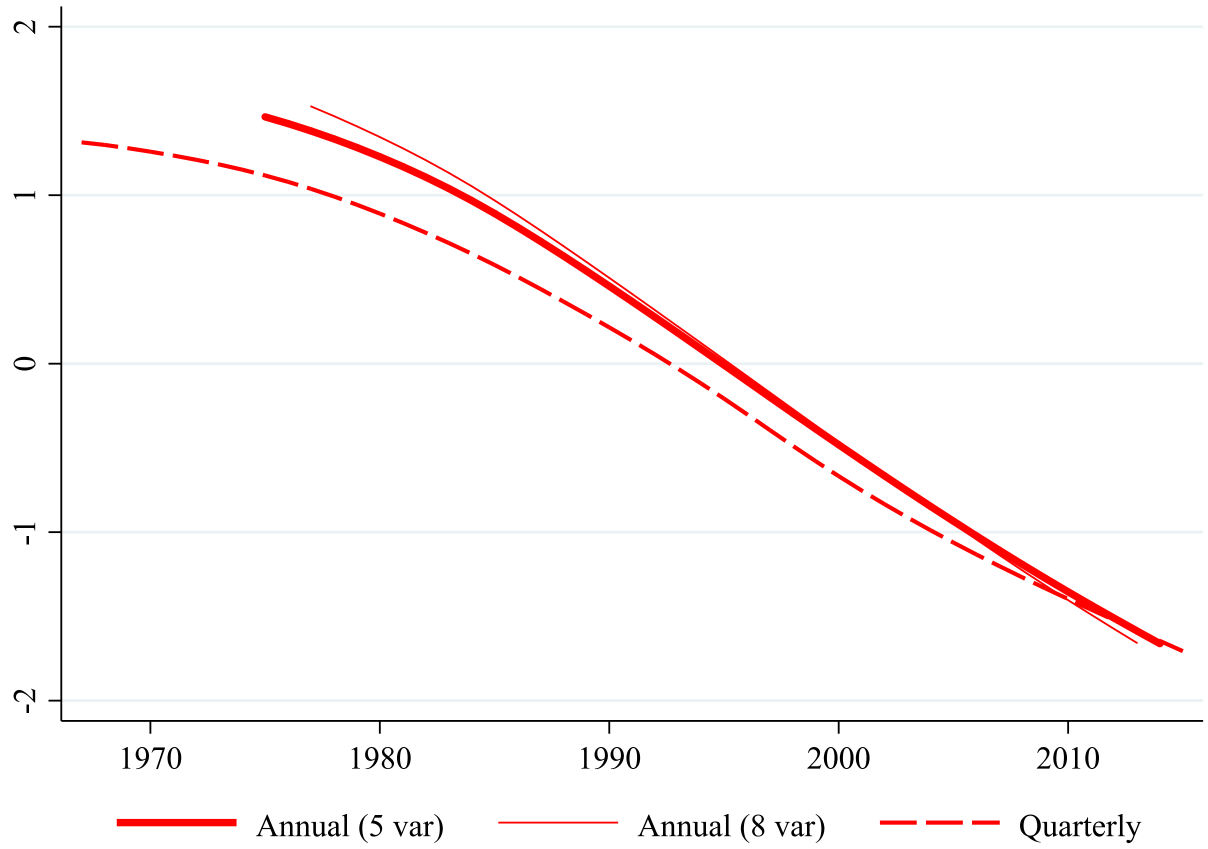


Note: The figure shows four measures of labor market fluidity (solid lines): from employment to unemployment (EU), from unemployment to employment (UE), from employment to out of the labor force (EN), and from out of the labor force to employment (NE). All series are recorded at quarterly frequency from 1967:Q2 to 2015:Q3 for a total of 194 observations. All flows are expressed as a ratio of persons in the first labor market state (e.g. EN shows the number of transitions from employment (E) to non employment (N) as a share of E). The dashed lines are the estimated low frequency trends using a biweight filter - bandwidth of 90 quarters. **Source:** authors' own calculation. **Raw data source:** data since 2012 are taken from Department of Labor, Bureau of Labor Statistics; data through 2012 are from Elsby et al. [25]. Their data is derived from three sources. From June 1967 to December 1975 the data are from Hoyt Bleakley (tabulations from Joe Ritter). From January 1976 through January 1990 the data are provided by Shimer [61]. From February 1990 onwards, the data are available from the Department of Labor, Bureau of Labor Statistics gross flow statistics.

Figure 2: Extracted trends - annual series.

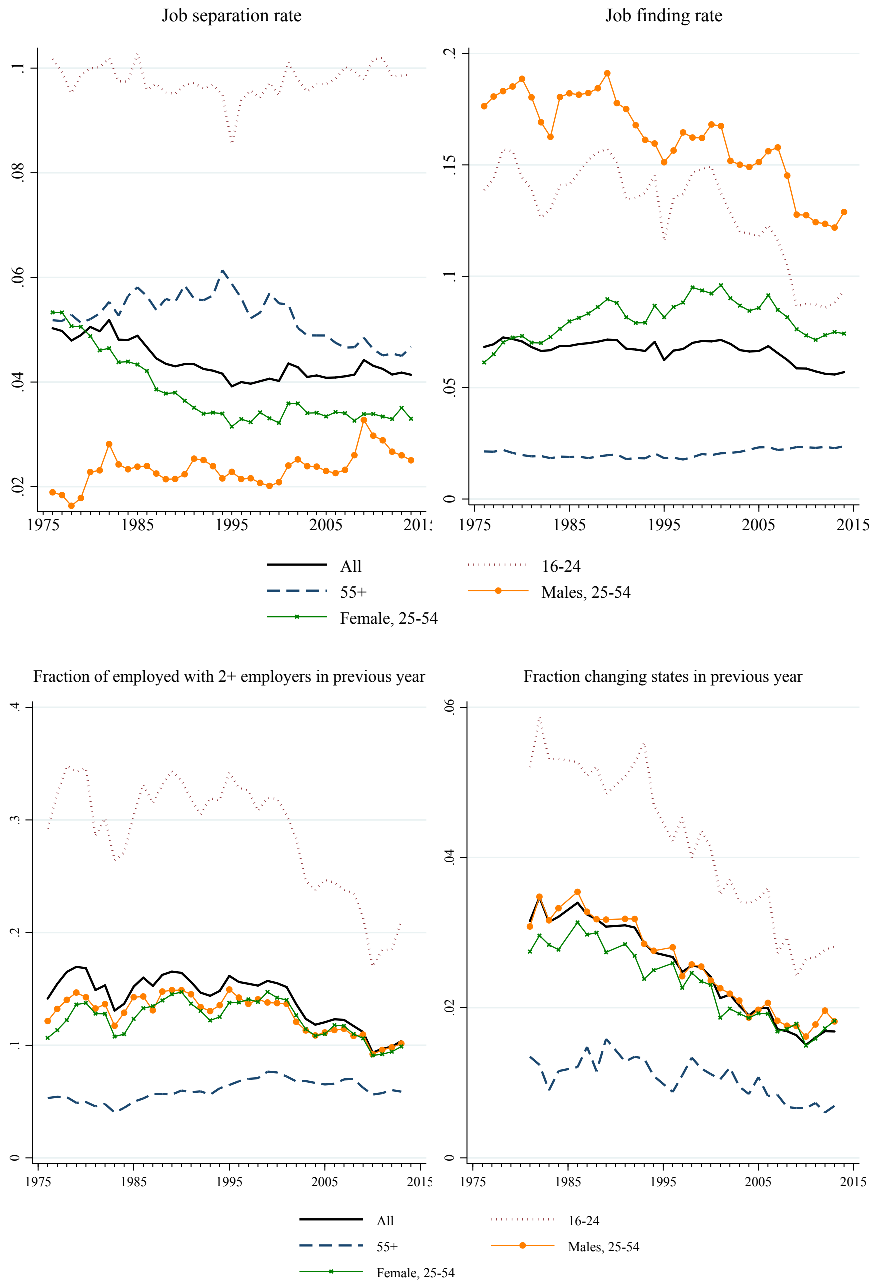
Note: The figure shows four measures of labor market fluidity (solid lines): the job-to-job (JtJ) transition rate (upper left panel), interstate migration (IM) rate (upper right panel), job destruction rate (JD, lower left panel), and job creation rate (JC, lower right panel). The dashed lines are the low frequency trends, estimated using a biweight filter - bandwidth of 30 years. All series are recoreded with annual frequency. EU, UE, EN, NE, JtJ and IM are recorded from 1975 to 2014 for a total of 40 observations. JDC and JC are recorded from 1977 to 2013 for a total of 37 observations. JtJ transition is expressed as a share of total employment. Interstate migration is expressed as a share of total population. JD and JC are expressed as ratios of average employment in the current period and the previous one. **Source:** microdata for the Current Population Survey Annual Social and Economic Supplement (CPS-ANES), as provided by the Unicon Research Corporation. Migration data come from the IRS Migration Data (https://www.irs.gov/uac/SOI-Tax-Stats-Migration-Data). Methodological changes make the post-2010 data not comparable to earlier years, so for years after 2010 we extend the IRS series with the growth rate of the migration rate from the U.S. Census Bureau, American Community Survey. For the years that the IRS and ACS data overlap, the level and changes in aggregate migration are quite similar (Molloy et al. [54]). JD and JC data are from Census Bureau, Business Dynamics Statistics (http://www.census.gov/ces/dataproducts/bds/index.html).

Figure 3: PCA - first component.



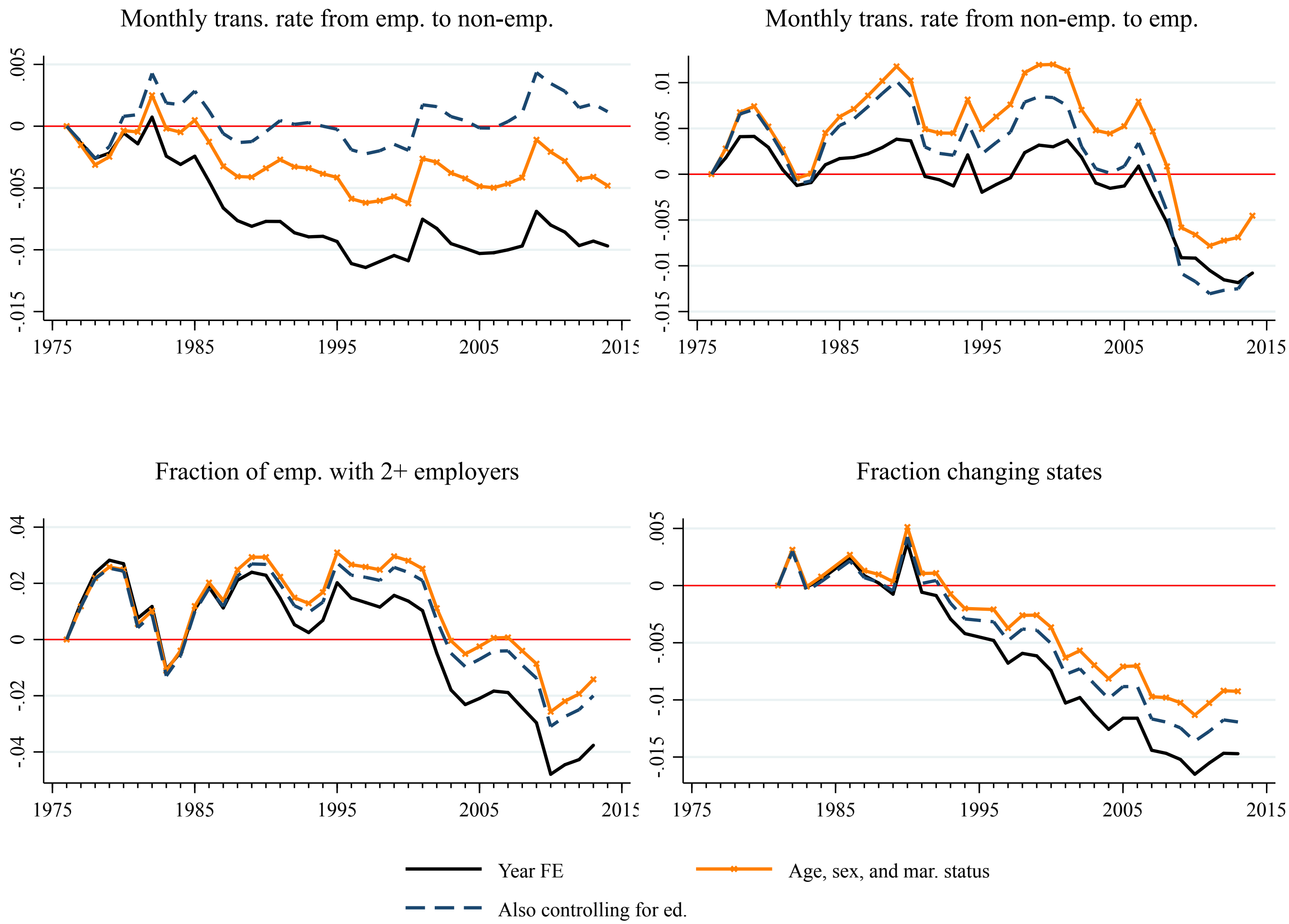
Note: The figure shows the (normalized) first component of a Principal Component Analysis (PCA). The PCA was run three times. The first time on four measures of labor market fluidity recorded at quarterly frequency (line “Quarterly” in the chart): from employment to unemployment (EU), from unemployment to employment (UE), from employment to out of the labor force (EN), and from out of the labor force to employment (NE). The second time the job-to-job transition (“JtJ”) rate - observed at annual frequency - was added (line “Annual (5 var)” in the chart). In this case the four quarterly flows have been transformed into annual observations first. Finally, the third time the PCA was run adding the interstate migration (IM) rate, job destruction (JD) and job creation (JC) rates (line “Annual (8 var)” in the chart), which are also observed at annual frequency. **Source:** authors’ own calculation. Raw data source for EU, UN, EN, and NE: data since 2012 are taken from Department of Labor, Bureau of Labor Statistics; data through 2012 are from Elsby et al. [25]. Their data is derived from three sources. From June 1967 to December 1975 the data are from Hoyt Bleakley (tabulations from Joe Ritter). From January 1976 through January 1990 the data are provided by Shimer (2012). From February 1990 onwards, the data are available from the BLS gross flow statistics. Raw data source for JtJ and IM: microdata for the Current Population Survey Annual Social and Economic Supplement (CPS-ANES), as provided by the Unicon Research Corporation. Migration data come from the IRS Migration Data (https://www.irs.gov/uac/SOI-Tax-Stats-Migration-Data). Methodological changes make the post-2010 data not comparable to earlier years, so for years after 2010 we extend the IRS series with the growth rate of the migration rate from the American Community Survey. For the years that the IRS and ACS data overlap, the level and changes in aggregate migration are quite similar (Molloy et al. [54]). JD and JC data are from U.S. Census Bureau, Business Dynamics Statistics (http://www.census.gov/ces/dataproducts/bds/index.html).

Figure 4: Job finding and separation rates by demographics - Job-to-job transitions and cross-state migration by demographics.



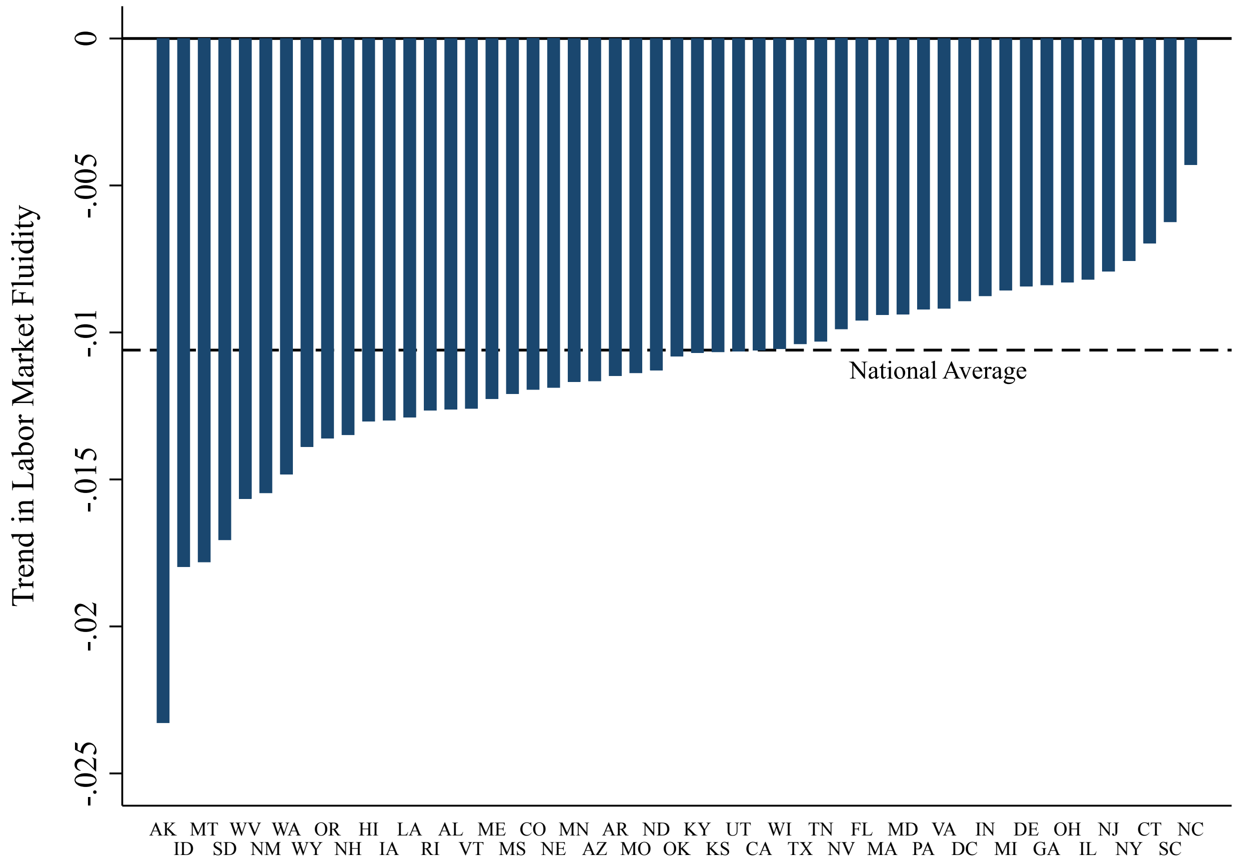
Note: Flows by demographic characteristics are estimated from matched CPS monthly data, authors' calculations. Estimates are derived from data from the March Supplement to the CPS. **Source:** authors' calculations.

Figure 5: Changes in flows after controlling for demographics.



Note: Figures plot the coefficients on year fixed effects after controlling for the listed demographic characteristics. All year fixed effects are shown relative to the 1976 estimate (or the 1981 estimate, for migration). **Source:** Authors' calculations..

Figure 6: Trend in labor market fluidity in each state.



Note: Trend in labor market fluidity is the first component from a Principal Component Analysis of linear trends of 8 annual variables: flows from employment to unemployment, flows from employment to out of labor force, flows from out of labor force to employment, flows from unemployment to employment, job to job changes, interstate migration, job creation and job destruction. Linear trends are estimated for each variable from a state-specific regression of the variable on a linear trend and the state unemployment rate (contemporaneous and 1 lag), estimated from 1980 to 2013. The national average is the weighted average of the state-specific trends, where the weights are the average state population from 1980 to 2013. **Source:** see text for data sources.

Figure 7: State-Level Correlation of Declines in Labor Market Fluidity of Older Workers and Youth Share.

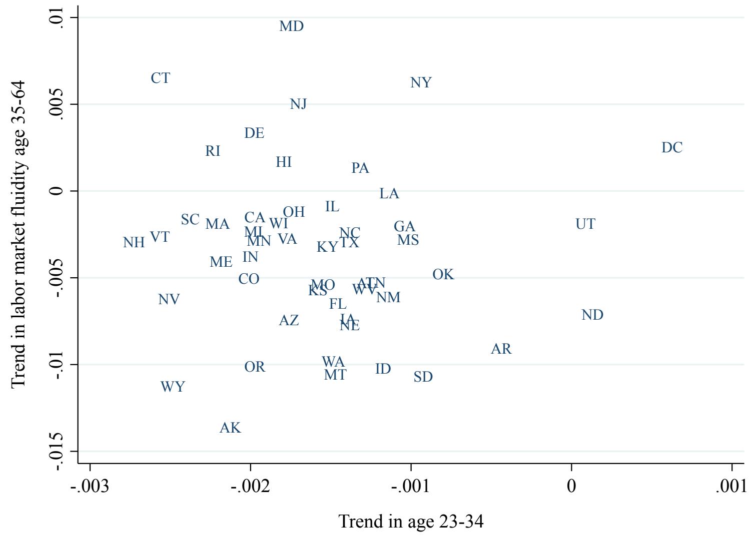
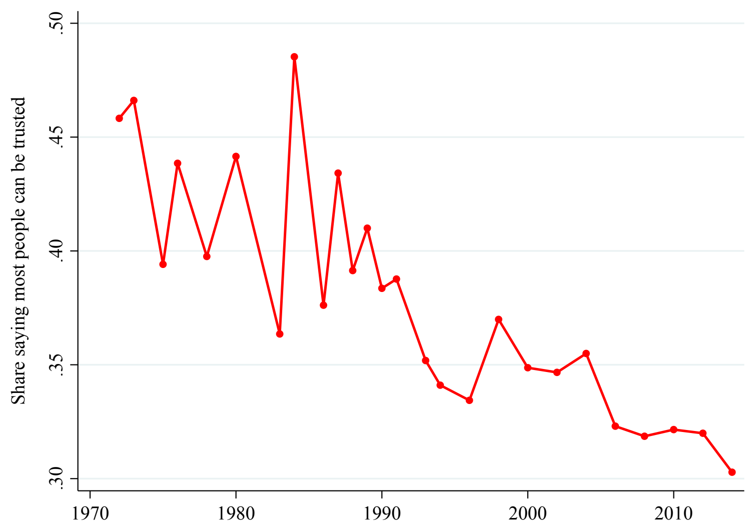


Figure 8: GSS trend: Social capital measure (trust) is declining.



Note: Trend in labor market fluidity is the first component from a Principal Component Analysis of linear trends of 8 annual variables: flows from employment to unemployment, flows from employment to out of labor force, flows from out of labor force to employment, flows from unemployment to employment, job to job changes, interstate migration, job creation and job destruction. Linear trends are estimated for each variable from a state-specific regression of the variable on a linear trend and the state unemployment rate (contemporaneous and 1 lag), estimated from 1998 to 2013. The trend in age is estimated from a state-specific regression of the fraction of the state's population age 23 to 34 on a linear time trend and the state unemployment rate (contemporaneous and 1 lag), estimated using annual data from 1980 to 2013. **Source:** authors' own calculations. see text for raw data sources.

Figure 9: GSS Scatter: States with bigger declines in trust had bigger declines in fluidity.

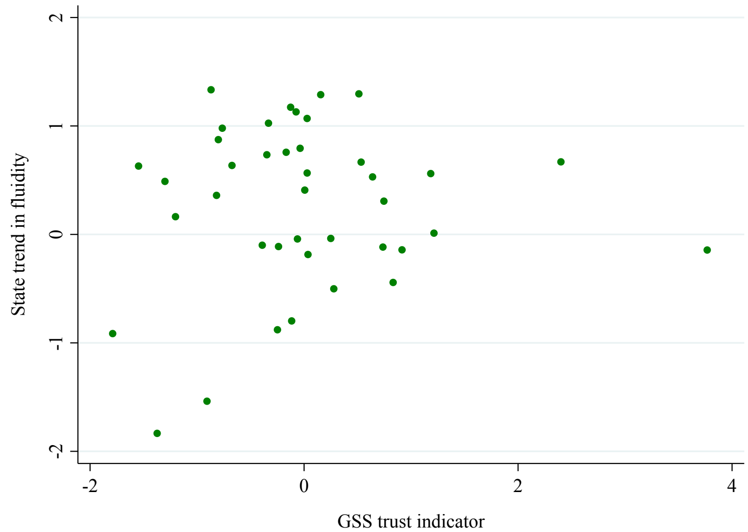
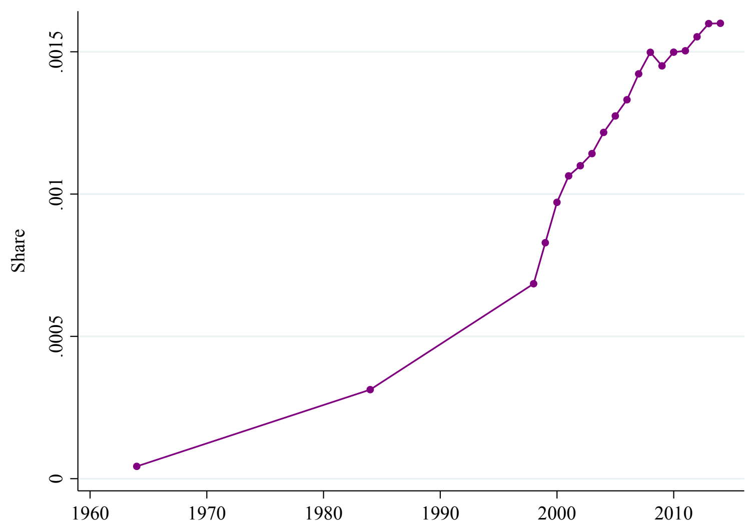


Figure 10: SHRM Members relative to US labor force.



Note: Trend in labor market fluidity is the first component from a Principal Component Analysis of linear trends of 8 annual variables: flows from employment to unemployment, flows from employment to out of labor force, flows from out of labor force to employment, flows from unemployment to employment, job to job changes, interstate migration, job creation and job destruction. Linear trends are estimated for each variable from a state-specific regression of the variable on a linear trend and the state unemployment rate (contemporaneous and 1 lag), estimated from 1980 to 2013. Figure 10 shows the share of SHRM members relative to the US labor force. **Source for Figure 9**: National Opinion Research Center, University of Chicago, General Social Survey (GSS) (sensitive data files) and Department of Labor, Bureau of Labor Statistics. f **Source for Figure 10:** 1998-2014 data on total SHRM members from Society for Human Resource Management personal communication. Member totals for 1964 and 1984 from SHRM website (accessed December 2015). US labor force totals from Department of Labor, Bureau of Labor Statistics.

Figure 11: State-Level Correlation of Declines in Labor Market Fluidity and Human Resource Membership.

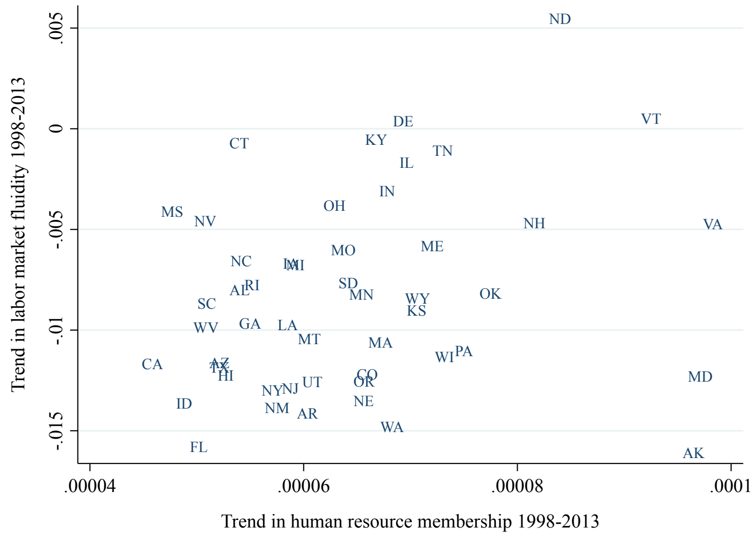
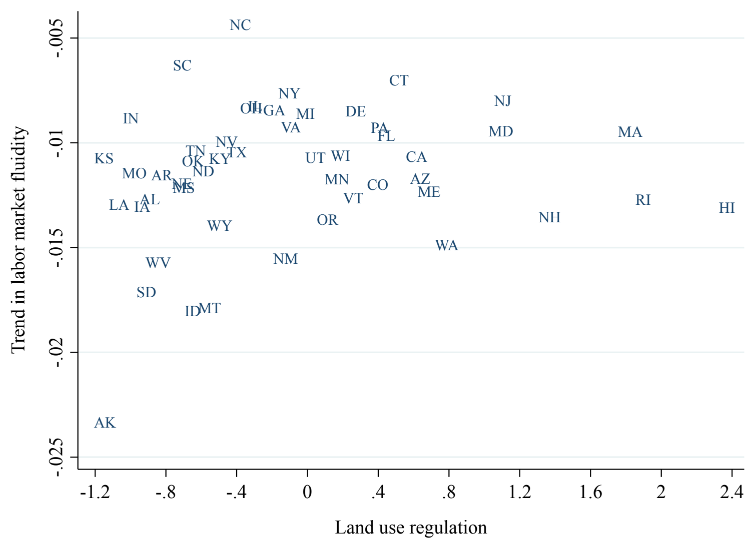


Figure 12: State-Level Correlation of Declines in Labor Market Fluidity and Land Use Regulation.



Note: Trend in labor market fluidity is the first component from a Principal Component Analysis of linear trends of 8 annual variables: flows from employment to unemployment, flows from employment to out of labor force, flows from out of labor force to employment, flows from unemployment to employment, job to job changes, interstate migration, job creation and job destruction. Linear trends are estimated for each variable from a state-specific regression of the variable on a linear trend and the state unemployment rate (contemporaneous and 1 lag), estimated from 1980 to 2013. Land use regulation is the state average of the Wharton Residential Land Use Regulation Index (Gyourko et al. [36]). AK and HI omitted in Figure 9 due to insufficient data for trend estimation. Regression coefficient (std error) is 0.053 (0.13) in figure with all regions. The trend in human resource membership is estimated from a state-specific regression of the fraction of the state's labor force who are members of the Society of Human Resource Managers on a linear time trend and the state unemployment rate (contemporaneous and 1 lag), estimated using annual data from 1998 to 2013. **Source:** authors' own calculations. **Raw data source** (for Figure 9): General Social Survey (sensitive data files) and Bureau of Labor Statistics.

Table 1: Descriptive Statistics of low frequency components.

			Minimum		Maximum	
	Mean	*Std. Dev.*	*Value*	*Year*	*Value*	*Year*
EU	.014	.001	.014	2014	.015	1978
UE	.259	.008	.242	2014	.271	1975
EN	.029	.001	.027	2014	.032	1975
NE	.047	.001	.047	2014	.047	1995
JtJ	.139	.010	.119	2014	.154	1975
IM	.026	.001	.024	2014	.028	1975
JD	.149	.004	.142	2013	.156	1977
JC	.169	.001	.154	2013	.183	1977

Note: The table shows summary statistics calculated for the low frequency components extracted using a biweght filter. All series are recorded at annual frequency. For comparability, the summary statistics have been calculated with 1975 as a starting point (although EU, UE, EN, and NE date back to 1968). For JD and JC the latest observation in the sample is 2013. EU, UE, EN, and NE are expressed as a ratio of persons in the first labor market state (e.g. EN shows the number of transitions from employment (E) to non employment (N) as a share of E). Job-to-job (JtJ) transition is expressed as a share of total employment. Interstate migration (IM) is expressed as a share of total population. JD and JC are expressed as ratios of average employment in the current period and the previous one. **Source:** authors' own calculation. **Raw data source:** for EU, UN, EN, and NE: data since 2012 are taken from Department of Labor, Bureau of Labor Statistics; data through 2012 are from Elsby et al. [25]. Their data is derived from three sources. From June 1967 to December 1975 the data are from Hoyt Bleakley (tabulations from Joe Ritter). From January 1976 through January 1990 the data are provided by Shimer [61]. From February 1990 onwards, the data are available from the BLS gross flow statistics. Raw data source for JtJ and IM: microdata for the Current Population Survey Annual Social and Economic Supplement (CPS-ANES), as provided by the Unicon Research Corporation. Migration data come from the IRS Migration Data (https://www.irs.gov/uac/SOI-Tax-Stats-Migration-Data). Methodological changes make the post-2010 data not comparable to earlier years, so for years after 2010 we extend the IRS series with the growth rate of the migration rate from the American Community Survey. For the years that the IRS and ACS data overlap, the level and changes in aggregate migration are quite similar (Molloy et al. [54]). JD and JC data are from U.S. Census Bureau, Business Dynamics Statistics (http://www.census.gov/ces/dataproducts/bds/index.html).

Table 2: Correlation of low frequency components.

	EU	**UE**	**EN**	**NE**	**JtJ**	**IM**	**JD**	**JC**
EU	1							
UE	.95 [0.00]	1						
EN	.82 [0.00]	.92 [0.00]	1					
NE	-.07 [0.61]	-.15 [0.28]	-.48 [0.00]	1				
JtJ	.99 [0.00]	.99 [0.00]	.93 [0.00]	.04 [0.76]	1			
IM	.97 [0.00]	.98 [0.00]	.97 [0.00]	-.09 [0.56]	.98 [0.00]	1		
JD	.99 [0.00]	.99 [0.00]	.94 [0.00]	.03 [0.83]	.99 [0.00]	.99 [0.00]	1	
JC	.98 [0.00]	.99 [0.00]	.96 [0.00]	-.02 [0.87]	.99 [0.00]	.98 [0.00]	.99 [0.00]	1

Note: The table shows the pairwise correlations of the low frequency components extracted using a biweght filter (with a window of 30 years). Significance levels are reported in squared brackets. All series are recorded at annual frequency. EU stands for "employmento to unemployment", "UE" stands for "unemployment to employment", "EN" stands for "employment to out of labor force, "NE" stands for out of the labor force to employment, "JtJ" stands for joj-to-job flow, "IM" stands for interstate migration, "JD" stands for job destruction, and "JC" stands for job creation. **Source:** authors' own calculation. **Raw data source:** for EU, UN, EN, and NE data since 2012 are taken from Department of Labor, Bureau of Labor Statistics; data through 2012 are from Elsby et al. [25]. Their data is derived from three sources. From June 1967 to December 1975 the data are from Hoyt Bleakley (tabulations from Joe Ritter). From January 1976 through January 1990 the data are provided by Shimer [61]. From February 1990 onwards, the data are available from the BLS gross flow statistics. Raw data source for JtJ and IM: microdata for the Current Population Survey Annual Social and Economic Supplement (CPS-ANES), as provided by the Unicon Research Corporation. Migration data come from the IRS Migration Data (https://www.irs.gov/uac/SOI-Tax-Stats-Migration-Data). Methodological changes make the post-2010 data not comparable to earlier years, so for years after 2010 we extend the IRS series with the growth rate of the migration rate from the American Community Survey. For the years that the IRS and ACS data overlap, the level and changes in aggregate migration are quite similar (Molloy et al. [54]). JD and JC data are from Census Bureau, Business Dynamics Statistics (http://www.census.gov/ces/dataproducts/bds/index.html).

Table 3: Principal Component Analysis (PCA).

	Eigenvalues			Eigenvector			
	(i)	(ii)	(iii)		(i)	(ii)	(iii)
Comp 1	2.08	3.87	6.88	EU	0.49	0.50	0.38
	[0.52]	[0.77]	[0.86]				
Comp 2	1.33	1.10	1.10	UE	0.62	0.50	0.38
	[0.33]	[0.22]	[0.13]				
Comp 3	0.57	0.01	0.01	EN	0.60	0.48	0.37
	[0.14]	[0.00]	[0.00]				
Comp 4	0.01	0.00	0.00	NE	-0.02	-0.01	-0.00
	[0.00]	[0.00]	[0.00]				
Comp 5	-	0.00	0.00	JtJ	-	0.50	0.38
	-	[0.00]	[0.00]				
Comp 6	-	-	0.00	IM	-	-	0.38
	-	-	[0.00]				
Comp 7	-	-	0.00	JD	-	-	0.38
	-	-	[0.00]				
Comp 8	-	-	0.00	JC	-	-	0.38
	-	-	[0.00]				

Note: The table shows the results of a Principal Component Analysis (PCA) run on the low frequency components of the series (estimated with a biweight filter). The PCA has been run three times, corresponding to the three columns shown in this table. The first time (column (i)) the PCA has been run on the four series recorded at quarterly frequency: EU (employmento to unemployment), UE (unemployment to employment), EN (employment to out of labor force), and NE (out of the labor force to employment). The second time (column (ii)), the PCA has been run on the four quarterly flows annualized adding job-to-job (JtJ) transition rate. Finally, the third time (column (iii)) the PCA was run on all eight series, therefore adding interstate migration rate (IM), job destruction rate (JD), and job creation rate (JC). The table reports the eigenvalues (left panel) together with the fraction of the total variance explained by each component (in squared brackets). The right hand side of the table shows the entries of the eigenvector asociated with the first component. **Source:** authors' own calculation. **Raw data source:** for EU, UN, EN, and NE data since 2012 are taken from Department of Labor, Bureau of Labor Statistics; data through 2012 are from Elsby et al. [25]. Their data is derived from three sources. From June 1967 to December 1975 the data are from Hoyt Bleakley (tabulations from Joe Ritter). From January 1976 through January 1990 the data are provided by Shimer [61]. From February 1990 onwards, the data are available from the BLS gross flow statistics. Raw data source for JtJ and IM: microdata for the Current Population Survey Annual Social and Economic Supplement (CPS-ANES), as provided by the Unicon Research Corporation. Migration data come from the IRS Migration Data (https://www.irs.gov/uac/SOI-Tax-Stats-Migration-Data). Methodological changes make the post-2010 data not comparable to earlier years, so for years after 2010 we extend the IRS series with the growth rate of the migration rate from the American Community Survey. For the years that the IRS and ACS data overlap, the level and changes in aggregate migration are quite similar (Molloy et al. [54]). JD and JC data are from U.S. Census Bureau, Business Dynamics Statistics (http://www.census.gov/ces/dataproducts/bds/index.html).

Table 4: Correlations of State-Level Trends in Labor Market Fluidity with Select State Characteristics.

	All States	**Excl. AK**
% Admin. support occ., 1977-79	0.29**	0.33**
	(0.07)	(0.06)
% Operator occupation, 1977-79	0.33**	0.38**
	(0.08)	(0.07)
% Union member, trend	-0.18**	-0.13*
	(0.07)	(0.06)
% Age 35-44, trend	0.34**	0.06
	(0.08)	(0.09)
Middle Atlantic	0.57	0.49*
	(0.28)	(0.24)
Mountain	-0.92**	-0.92**
	(0.21)	(0.17)
Pacific	-0.99**	-0.91**
	(0.25)	(0.21)
Constant	0.21*	0.26**
	(0.08)	(0.07)
# Obs.	51	50
Adj. R-squared	0.80	0.79

Note: Each column reports the results of regressing the trend in labor market fluidity in each state on the characteristics named in the rows. Trend in labor market fluidity is the first component from a Principal Component Analysis of linear trends of 8 annual variables: flows from employment to unemployment, flows from employment to out of labor force, flows from out of labor force to employment, flows from unemployment to employment, job to job changes, interstate migration, job creation and job destruction. Linear trends are estimated for each variable from a state-specific regression of the variable on a linear trend and the state unemployment rate (contemporaneous and 1 lag), estimated from 1980 to 2013. Average fraction of workers in administrative support occupations and operator occupations are from the CPS ASEC, as provided from the IPUMS and defined by the variable occ1990. Data on union membership are from Hirsch et. al (2001) and data on age distribution are from the CPS ASEC. Trends in union membership and age share are estimated from state-specific regressions of each variable on a linear trend and the state unemployment rate (contemporaneous and 1 lag), estimated using annual data from 1980 to 2013. Standard errors are reported in parentheses. * and ** indicate significance at the 5% and 1% levels, respectively. **Source:** see text for data sources.

Table 5: Average Hourly Wage on Jobs Held for Less than One Year for Men Ages 22-33.

NLS Cohort:	**NLS-YM**	**NLSY79**	**NLSY97**
Average Wage	10.5	11.6	10.6
	(0.42)	(0.56)	(0.48)
N obs	3,165	5,450	4,756
Average wage	11.2	12.2	10.3
in low-educ subsample	(0.50)	(0.62)	(0.53)
Average wage	13.0	15.2	14.6
in high-educ subsample	(0.67)	(0.97)	(0.69)
Observation years	1966-1971	1979-1994	2002-2013
	’73, ’75, ’76		
	’78, ’80, ’81		

Note: Average wages computed as the constant (standard error) in a regression of real wages on a control for the national unemployment rate, dummies for age, race and education, using the NLS sample indicated in the column heading. Sample is restricted to those with less than one year of tenure on their main job.

Table 6: Implied Returns to a Third Year of Employer Tenure for Men Ages 22-33.

NLS Cohort:	**NLS-YM**	**NLSY79**	**NLSY97**
Industry tenure	0.016**	0.015**	0.005
	(0.006)	(0.005)	(0.007)
Occupation tenure	0.015**	0.016***	0.012
	(0.006)	(0.004)	(0.007)
Employer tenure	-0.012	0.002	0.0005
	(0.006)	(0.006)	(0.008)
N obs	11,466	19,363	15,842
Observation years	1966-1971 ’73, ’75, ’76 ’78, ’80, ’81	1979-1994	2002-2013

Note: Cells show implied returns to three years of tenure in designated category, holding other characteristics constant. Returns are calculated from estimates of Equation 6, as discussed in text, using the NLS sample indicated in the column heading. *** indicates significance of level coefficient at the .1% level, ** at the 1% level, and * at the 5% level.

Table 7: Nested Tests of Contracting Models.

	PSID	**PSID**	**CPS**	**CPS**	**NLSY79**	**NLSY97**
Age	21-64	22-33	21-64	22-29	22-33	22-33
Years	1981-2013	1981-2013	1979-2010	1979- 2010	1979-1994	2002- 2013
UR[current]						
1980s	0.017**	0.013*	0.008**	0.022**		
	(0.003)	(0.005)	(0.002)	(0.006)	0.003	
					(0.003)	
1990s	0.005	0.010	0.028**	-0.060*		
	(0.003)	(0.007)	(0.009)	(0.025)		
2000s	-0.004	-0.013**	0.007*	0.000		-0.003
	(0.002)	(0.004)	(0.003)	(0.008)		(0.006)
UR[began]						
1980s	0.010	0.011	0.001	0.006		
	(0.006)	(0.007)	(0.003)	(0.008)	-0.000	
					(0.005)	
1990s	-0.001	-0.009	0.008	-0.007		-0.010
	(0.004)	(0.008)	(0.006)	(0.013)		(0.018)
2000s	-0.018**	-0.017	-0.010**	-0.041*		
	(0.005)	(0.010)	(0.003)	(0.016)		
UR[min]						
1980s	-0.044**	-0.035**	-0.008	-0.038**		
	(0.007)	(0.011)	(0.005)	(0.013)	-0.031**	
					(0.007)	
1990s	-0.060**	-0.045**	-0.049**	0.040		
	(0.007)	(0.013)	(0.011)	(0.025)		
2000s	0.002	0.010	-0.004	0.024		-0.017
	(0.007)	(0.012)	(0.007)	(0.020)		(0.027)
Observations	37,769	14,657	69,042	15,827	19,628	7,853

Note: Table reproduced from Molloy et al. [55]. UR[current] is national unemployment rate for all workers 16 and up in current survey year in NLSY. In PSID and CPS, UR[current] is national unemployment rate in previous calendar year. UR[min] is minimum of national unemployment rates from year job began to current survey year (in NLSY) and to past calendar year (in PSID and CPS). UR[began] is national unemployment rate in calendar year that job began (NLSY, CPS and PSID). All PSID and NLSY regressions include a quadric time trend, individual fixed effects, employer tenure, employer tenure squared, age and age squared. CPS regressions include a quadratic time trend, employer tenure and tenure squared, age and age squared, a dummy for having been married, for being non-white, dummies for educational status, industry, and region. In PSID and CPS, estimates by decade are estimated from a single regression with decade dummies and interactions of decade dummies with labor market conditions. NLSY79 results are shown in between the rows for the 1980s and 1990s because this sample spans both decades. Standard errors are clustered by individual in PSID and NLSY; standard errors in CPS are robust standard errors. ** indicates significance at the 1% level and * at the 5% level.

Table 8: Geographic Dimension of Rising Firm Level Inequality.

Census division	**Percent change in 90-10 Ratio of Establishment Compensation 1982-90 to 2007-15**
New England (CT, ME, MA, NH, RI, VT)	0.37
Middle Atlantic (NJ, NY, PA)	0.31
East North Central (IL, IN, MI, OH, WI)	0.07
West North Central (IA, KS, MN, MO, NE, ND, SD)	0.07
South Atlantic (DE, DC, FL, GA, MD, NC, SC, VA, WV)	0.40
East South Central (AL, KY, MI, TN)	-0.01
West South Central (AR,LA, OK, TX)	0.31
Mountain (AZ, CO, ID, MT, NV, NM, UT, WY)	0.14
Pacific (AK, CA, HI, OR, WA)	0.34

Note: Division level 90-10 ratios of average total compensation by establishment from National Compensation Survey data, unpublished statistics, prepared by the Department of Labor, Bureau of Labor Statistics.

Appendix

Appendix A. Further results and robustness check

Table A.1: Correlation of low frequency components (Christiano and Fitzgerald [17] filter extracted).

	EU	UE	EN	NE	JtJ	IM	JD	JC
EU	1							
UE	.02 [0.88]	1						
EN	-.18 [0.20]	.87 [0.00]	1					
NE	.23 [0.11]	.40 [0.00]	-.08 [0.57]	1				
JtJ	.96 [0.00]	.89 [0.00]	.50 [0.00]	.43 [0.00]	1			
IM	.84 [0.00]	.88 [0.00]	.80 [0.00]	.08 [0.62]	.91 [0.00]	1		
JD	.96 [0.00]	.93 [0.00]	.43 [0.01]	.49 [0.00]	.99 [0.00]	.87 [0.00]	1	
JC	.87 [0.00]	.89 [0.00]	.82 [0.00]	.04 [0.79]	.91 [0.00]	.99 [0.00]	.86 [0.00]	1

Note: The table shows the pairwise correlations of the low frequency components extracted using the lowpass version of the Christiano and Fitzgerald [17] filter (set to retrieve cycles longer than 30 years). Significance levels are reported in squared brackets. All series are recorded at annual frequency. EU stands for "employmento to unemployment", "UE" stands for "unemployment to employment", "EN" stands for "employment to out of labor force, "NE" stands for out of the labor force to employment, "JtJ" stands for joj-to-job flow, "IM" stands for interstate migration, "JD" stands for job destruction, and "JC" stands for job creation. **Source:** authors' own calculation. **Raw data source:** see main text for raw data source.

Table A.2: Correlation of low frequency components (cosine projections method).

	EU	UE	EN	NE	JtJ	IM	JD	JC
EU	1							
UE	.82 [0.88]	1						
EN	.05 [0.74]	.61 [0.00]	1					
NE	.99 [0.00]	.73 [0.00]	-.10 [0.49]	1				
JtJ	.99 [0.00]	.96 [0.00]	.34 [0.31]	.97 [0.00]	1			
IM	.76 [0.00]	.95 [0.00]	.79 [0.00]	.68 [0.00]	.84 [0.00]	1		
JD	.96 [0.00]	.99 [0.00]	.49 [0.00]	.93 [0.00]	.98 [0.00]	.93 [0.00]	1	
JC	.91 [0.00]	.99 [0.00]	.62 [0.00]	.87 [0.00]	.94 [0.00]	.97 [0.00]	.99 [0.00]	1

Note: The table shows the pairwise correlations of the low frequency components extracted using the cosines projetion method suggested by Müller and Watson [56] (on 2 cosine functions). Significance levels are reported in squared brackets. All series are recorded at annual frequency. EU stands for "employmento to unemployment", "UE" stands for "unemployment to employment", "EN" stands for "employment to out of labor force, "NE" stands for out of the labor force to employment, "JtJ" stands for joj-to-job flow, "IM" stands for interstate migration, "JD" stands for job destruction, and "JC" stands for job creation. **Source:** authors' own calculation. **Raw data source:** for EU, UN, EN, and NE data since 2012 are taken from Department of Labor, Bureau of Labor Statisticss (BLS); data through 2012 are from Elsby et al. [25]. Their data is derived from three sources. From June 1967 to December 1975 the data are from Hoyt Bleakley (tabulations from Joe Ritter). From January 1976 through January 1990 the data are provided by Shimer [61]. From February 1990 onwards, the data are available from the BLS gross flow statistics. Raw data source for JtJ and IM: microdata for the Current Population Survey Annual Social and Economic Supplement (CPS-ANES), as provided by the Unicon Research Corporation. Migration data come from the IRS Migration Data (https://www.irs.gov/uac/SOI-Tax-Stats-Migration-Data). Methodological changes make the post-2010 data not comparable to earlier years, so for years after 2010 we extend the IRS series with the growth rate of the migration rate from the American Community Survey. For the years that the IRS and ACS data overlap, the level and changes in aggregate migration are quite similar (Molloy et al. [54]). JD and JC data are from Census Bureau, Business Dynamics Statistics (http://www.census.gov/ces/dataproducts/bds/index.html).

Table A.3: Principal Component Analysis (PCA) - Christiano and Fitzgerald [17] lowpass filter.

	Eigenvalues			**Eigenvector**			
	(i)	(ii)	(iii)		(i)	(ii)	(iii)
Comp 1	2.24 [0.56]	3.25 [0.65]	6.08 [0.76]	EU	0.31	0.50	0.38
Comp 2	1.50 [0.37]	1.36 [0.27]	1.63 [0.20]	UE	0.58	0.53	0.38
Comp 3	0.24 [0.06]	0.37 [0.07]	0.27 [0.03]	EN	0.59	0.32	0.27
Comp 4	0.00 [0.00]	0.00 [0.00]	0.00 [0.00]	NE	0.45	0.24	0.10
Comp 5	- -	0.00 [0.00]	0.00 [0.00]	JtJ	-	0.54	0.39
Comp 6	- -	- -	0.00 [0.00]	IM	-	-	0.39
Comp 7	- -	- -	0.00 [0.00]	JD	-	-	0.39
Comp 8	- -	- -	0.00 [0.00]	JC	-	-	0.39

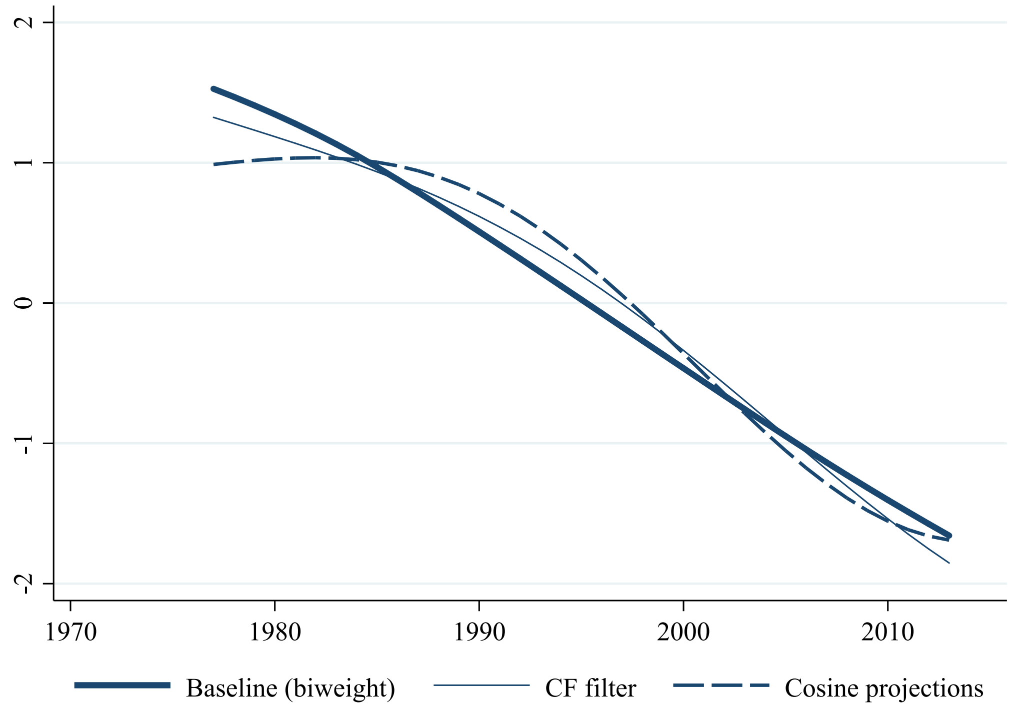
Note: The table shows the results of a Principal Component Analysis (PCA) run on the low frequency components of the series (estimated with the lowpass version of the Christiano and Fitzgerald [17] filter). The PCA has been run three times, corresponding to the three columns shown in this table. The first time (column (i)) the PCA has been run on the four series recorded at quarterly frequency: EU (employmento to unemployment), UE (unemployment to employment), EN (employment to out of labor force), and NE (out of the labor force to employment). The second time (column (ii)), the PCA has been run on the four quarterly flows annualized adding job-to-job (JtJ) transition rate. Finally, the third time (column (iii)) the PCA was run on all eight series, therefore adding interstate migration rate (IM), job destruction rate (JD), and job creation rate (JC). The table reports the eigenvalues (left panel) together with the fraction of the total variance explained by each component (in squared brackets). The right hand side of the table shows the entries of the eigenvector asociated with the first component. **Source:** authors' own calculation. **Raw data source:** for EU, UN, EN, and NE data since 2012 are taken from Department of Labor, Bureau of Labor Statistics (BLS); data through 2012 are from Elsby et al. [25]. Their data is derived from three sources. From June 1967 to December 1975 the data are from Hoyt Bleakley (tabulations from Joe Ritter). From January 1976 through January 1990 the data are provided by Shimer [61]. From February 1990 onwards, the data are available from the BLS gross flow statistics. Raw data source for JtJ and IM: microdata for the Current Population Survey Annual Social and Economic Supplement (CPS-ANES), as provided by the Unicon Research Corporation. Migration data come from the IRS Migration Data (https://www.irs.gov/uac/SOI-Tax-Stats-Migration-Data). Methodological changes make the post-2010 data not comparable to earlier years, so for years after 2010 we extend the IRS series with the growth rate of the migration rate from the American Community Survey. For the years that the IRS and ACS data overlap, the level and changes in aggregate migration are quite similar (Molloy et al. [54]). JD and JC data are from Census Bureau, Business Dynamics Statistics (http://www.census.gov/ces/dataproducts/bds/index.html).

Table A.4: Principal Component Analysis (PCA) - Müller and Watson [56] cosine projection method.

	Eigenvalues				Eigenvector		
	(i)	(ii)	(iii)		(i)	(ii)	(iii)
Comp 1	2.58	3.97	6.84	EU	0.58	0.49	0.36
	[0.64]	[0.79]	[0.85]				
Comp 2	1.41	1.01	1.14	UE	0.54	0.49	0.38
	[0.35]	[0.20]	[0.14]				
Comp 3	0.00	0.01	0.01	EN	0.15	0.19	0.20
	[0.00]	[0.00]	[0.00]				
Comp 4	0.00	0.00	0.00	NE	0.58	0.47	0.35
	[0.00]	[0.00]	[0.00]				
Comp 5	-	0.00	0.00	JtJ	-	0.49	0.37
	-	[0.00]	[0.00]				
Comp 6	-	-	0.00	IM	-	-	0.36
	-	-	[0.00]				
Comp 7	-	-	0.00	JD	-	-	0.38
	-	-	[0.00]				
Comp 8	-	-	0.00	JC	-	-	0.37
	-	-	[0.00]				

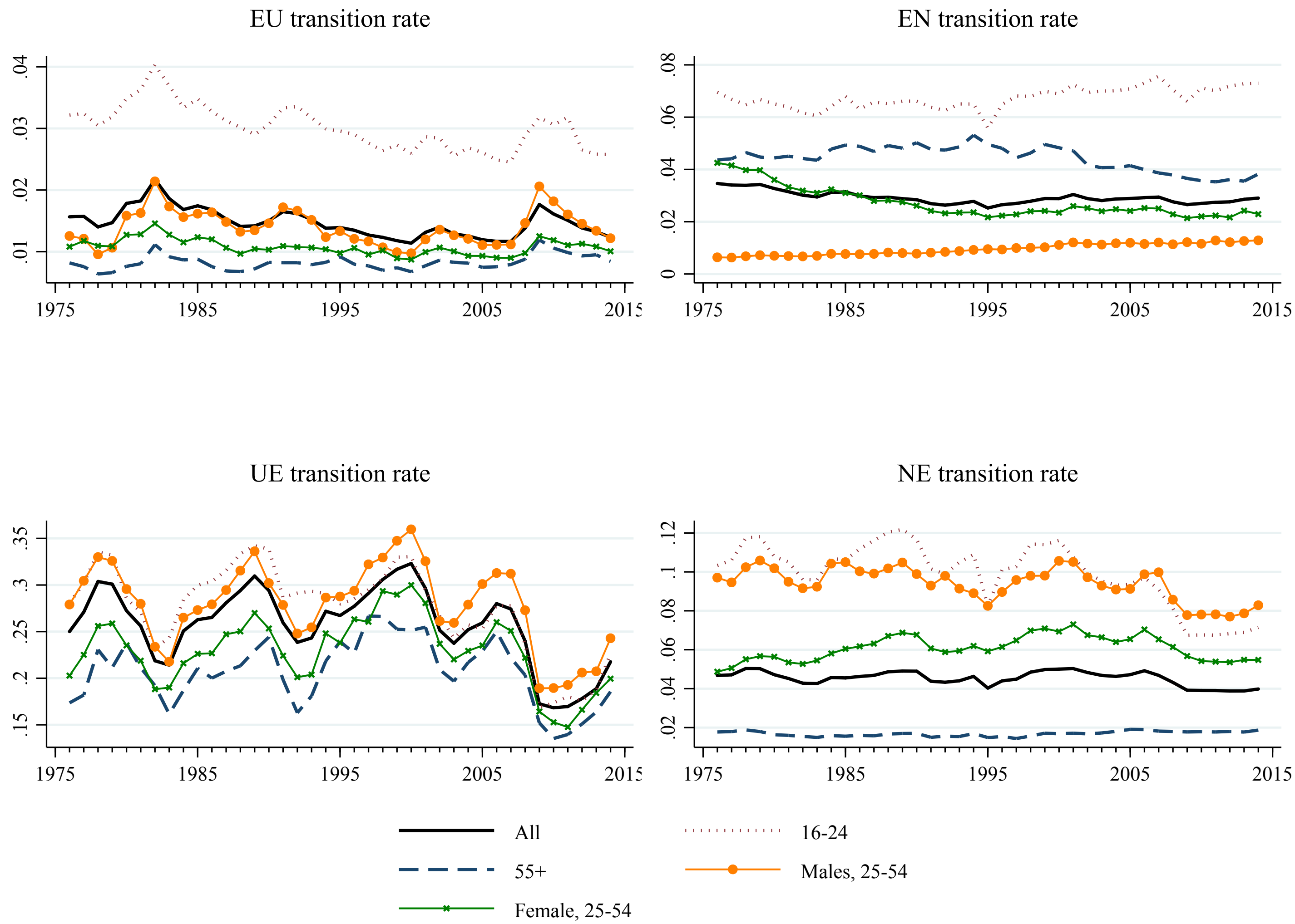
Note: The table shows the results of a Principal Component Analysis (PCA) run on the low frequency components of the series (estimated using the cosines projection method suggested by Müller and Watson [56]). The PCA has been run three times, corresponding to the three columns shown in this table. The first time (column (i)) the PCA has been run on the four series recorded at quarterly frequency: EU (employmento to unemployment), UE (unemployment to employment), EN (employment to out of labor force), and NE (out of the labor force to employment). The second time (column (ii)), the PCA has been run on the four quarterly flows annualized adding job-to-job (JtJ) transition rate. Finally, the third time (column (iii)) the PCA was run on all eight series, therefore adding interstate migration rate (IM), job destruction rate (JD), and job creation rate (JC). The table reports the eigenvalues (left panel) together with the fraction of the total variance explained by each component (in squared brackets). The right hand side of the table shows the entries of the eigenvector asociated with the first component. **Source:** authors' own calculation. **Raw data source:** for EU, UN, EN, and NE data since 2012 are taken from BLS; data through 2012 are from Elsby et al. [25]. Their data is derived from three sources. From June 1967 to December 1975 the data are from Hoyt Bleakley (tabulations from Joe Ritter). From January 1976 through January 1990 the data are provided by Shimer [61]. From February 1990 onwards, the data are available from the Department of Labor, Bureau of Labor Statistics (BLS) gross flow statistics. Raw data source for JtJ and IM: microdata for the Current Population Survey Annual Social and Economic Supplement (CPS-ANES), as provided by the Unicon Research Corporation. Migration data come from the IRS Migration Data (https://www.irs.gov/uac/SOI-Tax-Stats-Migration-Data). Methodological changes make the post-2010 data not comparable to earlier years, so for years after 2010 we extend the IRS series with the growth rate of the migration rate from the American Community Survey. For the years that the IRS and ACS data overlap, the level and changes in aggregate migration are quite similar (Molloy et al. [54]). JD and JC data are from Census Bureau, Business Dynamics Statistics (http://www.census.gov/ces/dataproducts/bds/index.html).

Figure A.1: PCA - first component - comparison.



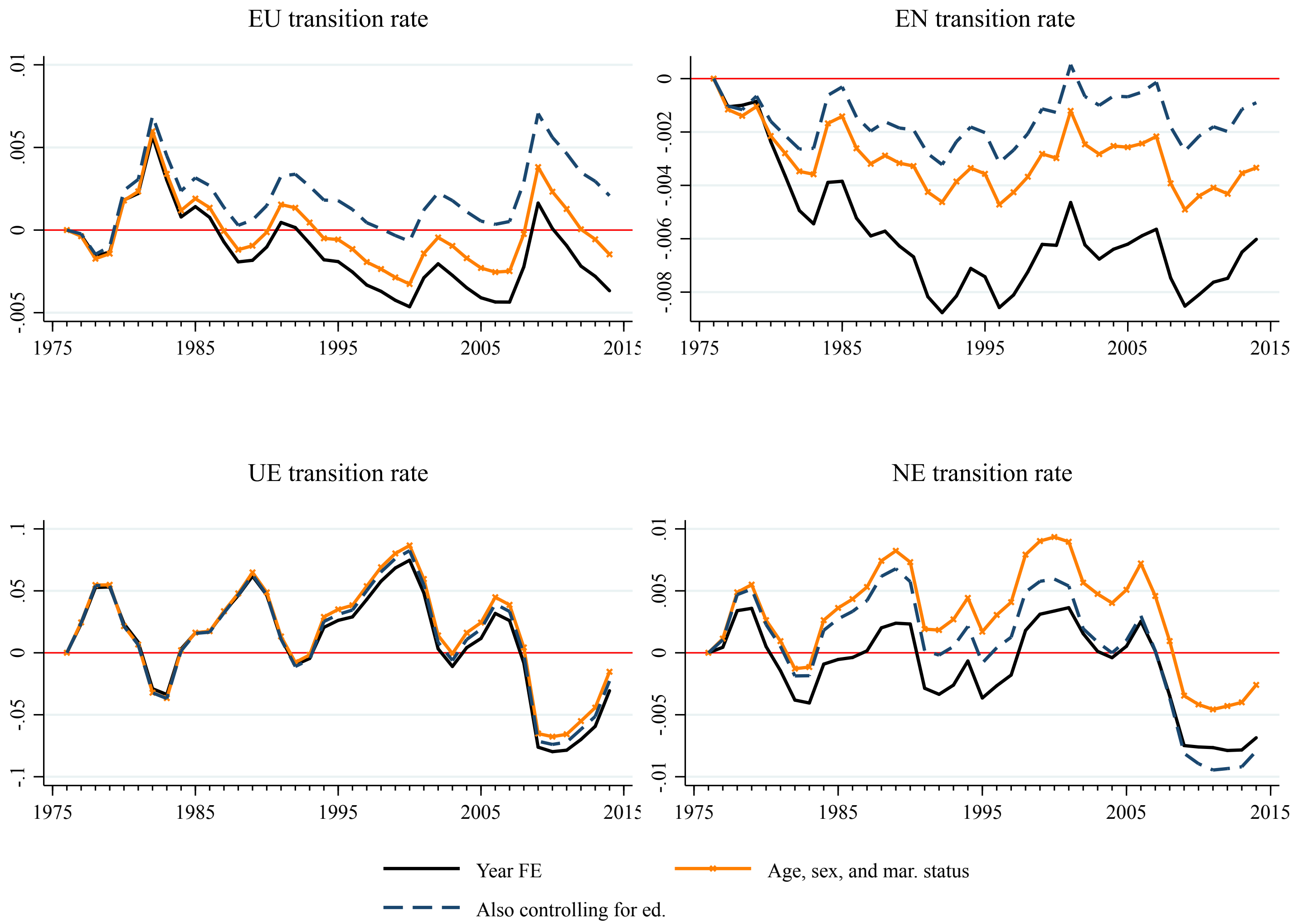
Note: The figure shows the first component(s) of a Principal Component Analysis. The PCA was run three times on the estimated low frequency component of eight labor market fluidity measures. The first time ("Baseline (biweight)"), the PCA was run on the trends estimated using a biweight filter (with a window of 30 years). The second time the PCA was run on the trends estimated using the lowpass version of the Christiano and Fitzgerald [17] filter (set to retreive cycles longer than 30 years). the third time the PCA was run on the trends estimated using the cosines projection method suggested by Müller and Watson [56]. In all cases the PCA was run on eight series recorded at annual frequency. The eight labor market fluidity measures included in the PCA are: swithcing rate from employment to unemployment (EU), from unemployment to employment (UE), from employment to out of the labor force (EN), and from out of the labor force to employment (NE), plus job-to-job transition ("JtJ") the Interstate Migration ("IM") rate, . job destruction ("JD"), and job creation ("JC") rates. EU, UE, EN, NE, JtJ and IM are recorded from 1975 to 2014 for a total of 40 observations. JDC and JC are recorded from 1977 to 2013 for a total of 37 observations. **Source:** authors' own calculation. See main text for raw data source.

Figure A.2: Labor market flows by demographics.



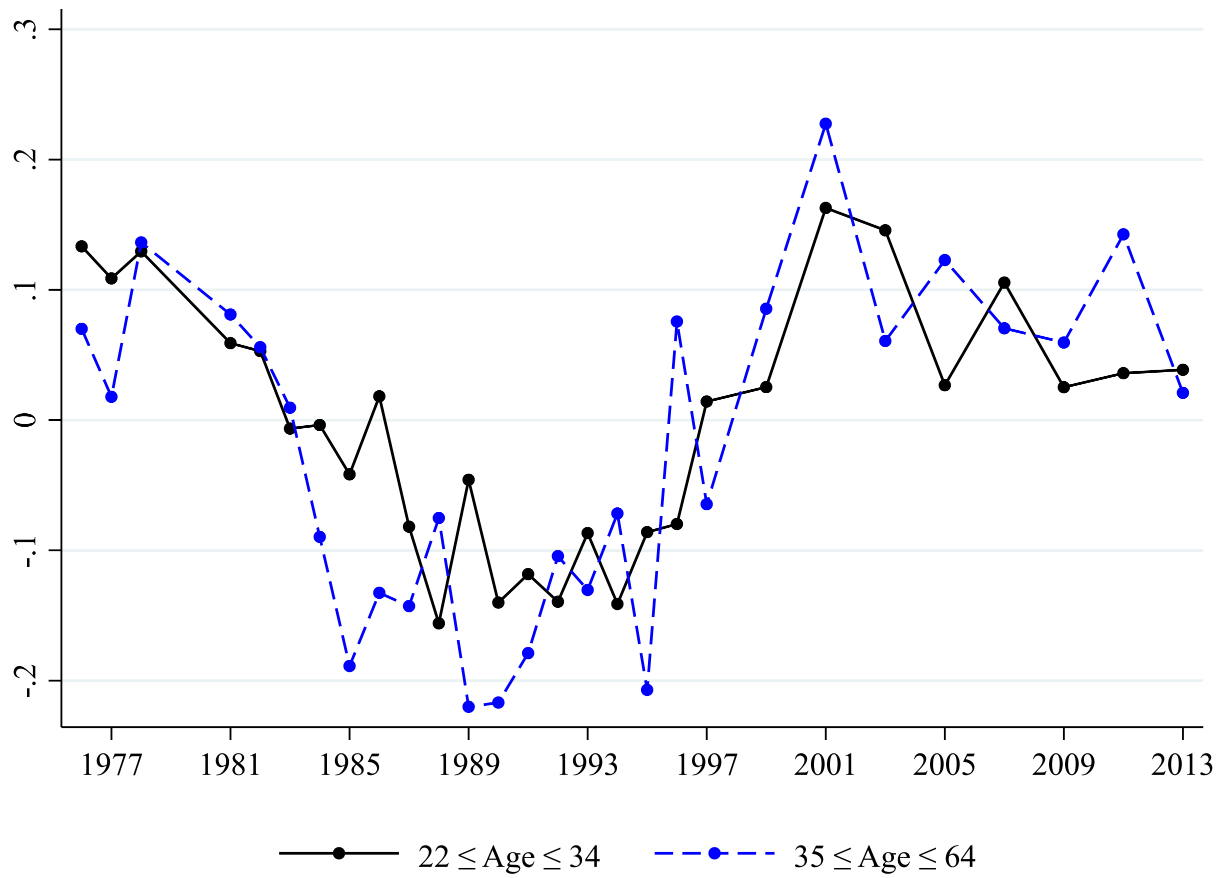
Note: Flows by demographic characteristics are estimated from matched Current Population Survey (CPS) monthly data, authors' calculations..

Figure A.3: Changes in flows after controlling for demographics.



Note: Figures plot the coefficients on year fixed effects after controlling for the listed demographic characteristics. All year fixed effects are shown relative to the 1976 estimate (or the 1981 estimate, for migration). Flows by demographic characteristics are estimated from matched Current Population Survey (CPS) data. Authors' calculations..

Figure A.4: Average Starting Wages in the PSID.



Note: Figure shows the average residual from a regression of the log of real wages among men with less than 12 months of tenure. The regression includes indicators for age group (30 -39, 40-49 and 50-64), race (black, Hispanic, other race), educational attainment (less than high school, some college, college or more) and the national unemployment rate. Regression is estimated separately for men age 22 to 34 and men age 35 to 64. Wages are the hourly wage for hourly workers and salary divided by usual hours for salaried workers, and are deflated by the price index for personal consumption expenditures. Workers earning less than half of the federal minimum wage are excluded. **Source:** author's calculations from the University of Michigan, Panel Study of Income Dynamics (PSID).

Table A.5: Correlations of State-Level Trends in Labor Market Fluidity with State Characteristics.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
% Age 18-22, 1977-79	0.28								
% Age 23-34, 1977-79	0.01								
% Age 35-44, 1977-79	0.31								
% Age 45-64, 1977-79	0.64*								
% Age 64+, 1977-79	-0.07								
% Age 18-22, trend	0.27								
% Age 23-34, trend	0.22								
% Age 35-44, trend	0.63*								
% Age 45-64, trend	0.10								
% Age 64+, trend	-0.01								
% Less than High School, 1977-79		0.31							
% Some College, 1977-79		-0.43							
% College Plus, 1977-79		-0.16							
% Less than High School, trend		0.15							
% Some College, trend		0.07							
% College Plus, trend		0.75*							
% Married, 1977-79			-0.72						
% Married, trend			0.07						
% Single (never married), 1977-79			-0.35						
% Single (never married), trend			-0.21						
% Homeowner, 1977-79				-0.06					
% Homeowner, trend				-0.09					
% Manufacturing, 1977-79					0.75*				
% Retail, 1977-79					-0.02				
% FIRE, 1977-79					0.11				
% Service, 1977-79					0.11				
% Agriculture, 1977-79					-0.51				
% Manufacturing, trend					0.31				
% Retail, trend					0.04				
% FIRE, trend					0.19				
% Service, trend					0.19				
% Agriculture, 1977-79					-0.28				

Table A.5 continues on the next page.

Table A.5 continued

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
% Manag./prof. occ., 1977-79						-0.05			
% Technicians, 1977-79						-1.69*			
% Service occupation, 1977-79						-0.08			
% Production/craft occupation, 1977						-0.14			
% Operator occupation, 1977-79						0.76			
% Sales occupation, 1977-79						0.78*			
% Admin. support occ., 1977-79						1.82*			
% Manag./prof. occ., trend						0.38*			
% Technicians, trend						-1.10			
% Service occupation, trend						0.08			
% Production/craft occupation, trend						-0.20			
% Operator occupation, trend						0.28			
% Sales occupation, trend						0.44			
% Admin. support occ., trend						1.21			
% Union member, 1977-79							-0.17		
% Union member, trend							-0.37		
% Self Employed, 1977-79								-0.39*	
% Self Employed, trend								0.01	
% Government worker, 1977-79								-0.49*	
% Government worker, trend								-0.15	
Northeast									0.37
Middle Atlantic									1.18*
East North Central									0.91*
West North Central									0.02
South Atlantic									0.49*
East South Central									0.40
West South Central									-0.20
Mountain									-1.12*
Pacific									-1.34*
# Obs.	51	51	51	51	51	51	51	51	51
Adj. R-squared	0.54	0.47	0.09	-0.03	0.43	0.56	0.04	0.31	0.54

Note: Each column reports the results of regressing the trend in labor market fluidity in each state on the characteristics named in the rows. Trend in labor market fluidity is the first component from a Principal Component Analysis of linear trends of 8 annual variables: flows from employment to unemployment, flows from employment to out of labor force, flows from out of labor force to employment, flows from unemployment to employment, job to job changes, interstate migration, job creation and job destruction. Linear trends are estimated for each variable from a state-specific regression of the variable on a linear trend and the state unemployment rate (contemporaneous and 1 lag), estimated from 1980 to 2013. All independent variables except union share are from the CPS ASEC, as provided from the IPUMS. Union membership is from Hirsch et al. [40]. Trends in independent variables are estimated from state-specific regressions of each variable on a linear trend and the state unemployment rate (contemporaneous and 1 lag), estimated using annual data from 1980 to 2013. Standard errors are reported in parentheses. * and ** indicate significance at the 5% and 1% levels, respectively. **Source:** see text for data sources.