The Great Increase in Relative Volatility of Real Wages in the United States*

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Abstract

This paper documents that over the past 25 years, real average hourly wages in the United States have become substantially more volatile relative to output. We use micro-data from the Current Population Survey (CPS) to show that this increase in relative volatility is predominantly due to increases in the relative volatility of hourly wages across different groups of workers. Compositional changes, on the other hand, account for at most 13% of the increase in relative wage volatility. Using a Dynamic Stochastic General Equilibrium (DSGE) model, we show that the observed increase in relative wage volatility is unlikely to come from changes outside of the labor market (e.g. smaller exogenous shocks or more aggressive monetary policy). By contrast, greater flexibility in wage setting due to deunionization and a shift towards performance-pay contracts as experienced by the U.S. labor market is capable of accounting for a substantial fraction of the observed increase in relative wage volatility. Greater wage flexibility also decreases the magnitude of business cycle fluctuations, suggesting an interesting new explanation for the Great Moderation.

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1 Introduction

The 25 years prior to the most recent recession were a time of unprecedented macroeconomic stability for the United States. During that period, referred to by many as the 'Great Moderation', the business cycle volatility of output declined by more than 50% and the volatility of many other macroeconomic aggregates fell by similar proportions.[1]

In this paper, we show that the Great Moderation does not apply to one of the most prominent labor market aggregates: real average hourly wages (or 'hourly wages' for short). Specifically, we document the following results:

1. From 1953-1984 to 1984-2006, the business cycle volatility of hourly wages increased between 15% and 60%, depending on the dataset and filtering method used.

2. As a result, the business cycle volatility of the aggregate wage *relative* to the volatility of aggregate output became about 2.5 to 3.5 times larger over the two sample periods.

The increase in volatility of hourly wages raises several questions. First, to what extent does this increase apply to different groups of workers? Second and related, how much of the increase in volatility is due to compositional changes of the workforce; i.e. a shift of the workforce towards jobs with more volatile wages? Third, to what extent is the increase in volatility related to structural changes in the U.S. labor market? Fourth, how do such labor market changes contribute to our understanding of business cycle fluctuations in general and the Great Moderation in particular?

To answer the first and second question, we use microdata from the Current Population Survey (CPS) to construct hourly wage series for different groups of workers. We document that the increase in *absolute* volatility of hourly wages is not generalized but concentrated among male, skilled and salaried workers. Also, there are large differences across industries, with absolute volatilities of hourly wages in some industries decreasing. However, these decreases are generally modest and thus, the volatility of hourly wages *relative* to the volatility of output increases substantially for all worker groups considered. We call this phenomenon the 'Great Increase in Relative Volatility of Real Wages'.

To quantify how much of the increase in the relative volatility of average hourly wages is due to compositional changes of the workforce towards jobs with more volatile wages, we develop an accounting method that allows us to decompose the increase in relative volatility of average hourly wages into different sources. The main result coming out of this exercise is that the widespread increase of relative wage volatility across different worker groups explains 70% or more of the

[1] See McConnell and Perez-Quiros (2000), Blanchard and Simon (2001) or Stock and Watson (2002).

increase in the relative volatility of average hourly wages. Compositional changes of the workforce, by contrast, account for at most 13% of the increase in relative wage volatility. This suggests that the increase in relative wage volatility is due to structural changes in the economic environment that affect wage dynamics of different workers in similar ways although to varying degrees.

To address the third and fourth question, we build a Dynamic Stochastic General Equilibrium (DSGE) model that allows us to assess the quantitative effects of deunionization and increased incidence of performance-pay. Our focus on these two particular structural changes is motivated by a combination of empirical observations. First, over the past decades, the U.S. labor market experienced a marked decline in private-sector unionization (e.g. Farber and Western, 2001) and a shift towards performance-pay contracts (e.g. Lemieux et al., 2009a). Second, Lemieux et al. (2009b) show that wages of non-union workers with performance-pay contracts are most responsive to local labor market shocks and least responsive for union workers without performance-pay. Exactly the opposite is the case for hours worked. Taken together, these observations suggest that deunionization and increased incidence of performance-pay result in greater wage flexibility, making wages more and hours (and output) less responsive to business cycle shocks.

We calibrate the model consistent with U.S. data to assess this conjecture. First, we show that while changes in the (absolute and relative) importance of exogenous shock processes can have a sizable effect on the *absolute* volatility and cyclicality of wages, their effect on the *relative* volatility of wages is negligible. Similarly, structural changes to the economy that do not directly affect the labor market (e.g. a more aggressive monetary policy response to inflation) are unlikely to have a large effect on the relative volatility of wages. We then decrease the proportion of unionized workers and increase the incidence of performance-pay contracts as observed in U.S. data. We find that these changes indeed lead to greater wage flexibility in equilibrium, accounting for a substantial fraction of the observed increase in relative wage volatility while simultaneously decreasing the magnitude of business cycle fluctuations. This suggests that the decline of unionization and the shift towards performance-pay contracts experienced by the U.S. labor market are at least partially responsible for the observed changes in wage dynamics and the Great Moderation of macroeconomic fluctuations.[2]

Our paper contributes to a recent literature on changes in U.S. labor market dynamics. Most notably, Gali and Gambetti (2009) and Stiroh (2009) document that the Great Moderation period is characterized by an increase in the relative volatility of hours worked and a fall in the correlation of labor productivity with output and hours. Gali and Van Rens (2010) build a DSGE model

[2] Increased wage flexibility does not render the economy immune to large business cycle shocks such as the ones experienced during the recent financial crisis. Our results suggest that the effects of these large shocks would have been more severe if wage setting had been as rigid as in the early 1980s.

with labor hoarding and argue that a decrease in labor hoarding due to decreased hiring costs accounts for both of these changes in labor market dynamics.[3] Gali and Van Rens (2010) also note the increase in relative wage volatility and argue that under certain assumptions about wage setting, a decrease in hiring costs increases wage flexibility.[4] Nucci and Riggi (2010) propose an alternative DSGE model where workers get paid separately for hours worked and effort, with the latter component being interpreted as performance-pay. Nucci and Riggi (2010) translate a shift towards performance-pay contracts as an increased sensitivity of the effort component of wages to current economic conditions and argue that this shift is capable of accounting for the empiricial evidence in Gali and Gambetti (2009). Compared to these papers, our paper focuses more squarely on wage volatility. In particular, we are the first to carefully analyze the characteristics of different aggregate wage series and document that the increase in relative wage volatility is not due to compositional changes of the workforce but generalized across different worker groups. This result is important because it directs our search for possible explanations towards structural changes in the economic environment that have a similar impact on all workers. Furthermore, our structural analysis of the general equilibrium effects of deunionization and increased incidence of performance-pay is based on a very different model of wage setting that can be calibrated explicitly with data on union density and the proportion of workers on performance-pay contracts. We are therefore capable to provide a *quantitative* assessment of the effects of structural changes in the U.S. labor market.[5]

The rest of the paper proceeds as follows. In Section 2 we document the increase in volatility of different aggregate hourly wage measures. Section 3 presents changes in relative wage volatility across different worker decompositions and implements the volatility accounting exercise. Section 4 describes our DSGE model and simulates the effects of deunionization and increased incidence of performance-pay. Section 5 concludes.

[3] In related work, Barnichon (2010) documents that the correlation of labor productivity with unemployment has switched from mildly negative to significantly positive during the Great Moderation. He proposes a combination of changes in the relative importance of business cycle shocks and lower labor search friction as a potential explanation.

[4] To our knowledge, the increase in relative wage volatility was first observed in unpublished manuscripts by Champagne (2007) and Gourio (2007). The results in Champagne (2007) provide the starting point for our paper.

[5] A number of other recent papers conjecture that different structural changes in the U.S. labor market have led to greater wage flexibility. Prominent examples are Blanchard and Gali (2007); Davis and Kahn (2008) or Lemieux et al. (2009a, b). Davis and Kahn (2008) conclude that greater wage flexibility *"...offers a unified explanation for the rise in wage and earnings inequality, flat or rising volatility in household consumption, a decline in the job-loss rate, and declines in firm-level and aggregate volatility measures."* However, none of these papers proposes a structural model that would allow an explicit *quantitative* assessment of these effects as is done in our paper.

2 Hourly wages during the Great Moderation

In this section, we document the increase in volatility of average real hourly wages in the United States. We first describe the construction of our preferred measure of hourly wages and present the main results. Then, we discuss robustness with respect to alternative hourly wage measures. For the sake of brevity, we keep the description of the data to a minimum. An extensive appendix, available on the authors' websites, provides more detailed information.

2.1 Data

The most comprehensive measure of average hourly wages in the non-farm business sector comes from the Bureau of Labor Statistics' (BLS) Labor Productivity and Costs (LPC) program. The measure is computed as total compensation based on the Quarterly Census of Employment and Wages (QCEW) divided by a corresponding series of total hours worked and is available quarterly starting in 1948.[6] The QCEW is a mandatory employer-based program for all employees covered by unemployment insurance (UI) and spans about 98% of U.S. establishments and jobs. Its total compensation measure includes direct wage and salary payments (including executive compensation); commissions, tips and bonuses; as well as supplements such as vacation pay or employer contributions to pension and health plans. To obtain real hourly wages, we deflate the LPC measure by the Personal Consumption Expenditure (PCE) index from the National Income and Products Accounts (NIPA) and check that the results are robust to other deflators. To compare our wage series with the business cycle, we use non-farm business real chain-weighted GDP per capita, obtained from the NIPA tables. All series are logged and filtered to extract the business cycle component. We use three different filtering methods: (i) quarterly first-difference filtering; (ii) Hodrick-Prescott (HP) filtering; and (iii) Bandpass Filtering (BP) as proposed by Christiano and Fitzgerald (2003).

2.2 Main results

Table 1 shows the standard deviation of output and real hourly wages for the periods 1953:2-1983:4 and 1984:1-2006:4, with standard errors for each estimate provided in brackets.[7] The sample split

[6]We restrict the analysis to the non-farm business sector because total economy-wide hours (and therefore economy-wide average hourly wages) are difficult to obtain on a quarterly basis for a sufficiently long sample and because it is unclear how to interpret public-sector wages in a market-based economy such as the one presented in Section 4. Including farming would not change any of the results.

[7]We start the sample in 1953:2 to avoid the extreme swings in inflation during the Korean War. Starting the sample in 1948 does not change any of the results. Standard errors are computed via the delta method from GMM-based estimates. See the appendix for details.

is motivated by the Great Moderation literature that estimates a break in output volatility in 1984 (e.g. McConnell and Perez-Quiros, 2000). While output volatility decreases by about 50% over the two periods (i.e. the Great Moderation), the volatility of hourly wages increases by 40% to 60% depending on the filtering method.[8] The p-value of Levene's (1960) test of equal variance indicates that this increase in hourly wage volatility is highly significant. The different evolution of output and hourly wage volatility is even more striking when considering relative standard deviations. As the last column of Table 1 shows, the volatility of hourly wages relative to the volatility of output increases more than three-fold over the two periods.

To further illustrate this result, we plot the volatility of output and hourly over 8-year rolling windows. As the first panel of Figure 1 illustrates, the volatility of output falls precipitously in the 1980s whereas the volatility of hourly wages increases during the 1980s and 1990s. As shown in the second panel, the relative volatility of hourly wages thus increases dramatically from the mid-1980s to the mid-1990s. Thereafter, the relative volatility of hourly wages returns to an intermediate level that remains, however, more than twice as high as the level before the mid-1980s.

We take away two main results from Table 1 and Figure 1. First, as the volatility of output drops during the Great Moderation, the *absolute* volatility of hourly wages increases. Second, the drop in output volatility is much larger in absolute terms than the increase in hourly wage volatility. The more than three-fold increase in the *relative* volatility of hourly wages is therefore driven to a large part by the drop in output volatility. The challenge for any theory is to explain how there can be such a marked fall in output volatility without a similar fall in wage volatility.

2.3 Evidence from alternative hourly wage measures

The hourly wage from the LPC is based on a very broad measure of compensation that includes employer contributions to pension and health plans as well as gains from exercising certain stock options. Both of these components have grown importantly over the past decades, raising the question of whether they drive the documented increase in hourly wage volatility. The question is particularly relevant for stock options because they are likely to be exercised in upturns when their value is higher than their fair-market value at the time they were granted (i.e. the time when they should have been recorded as compensation).[9] More generally, we may be concerned whether the

[8] Interestingly, if we used the GDP price index to deflate the LPC measure, the increase in hourly wage volatility would be even larger. See appendix.

[9] Mehran and Tracy (2001) argue that the increased incidence of stock options in the 1990s and their inclusion in compensation at the time of exercise instead of the time of grant has biased the evolution of compensation upwards. The authors also conjecture that increased use of stock options may render compensation more variable.

increase in hourly wage volatility is representative of the entire workforce or due to the emergence of a small fraction of high-earning individuals with very volatile compensation.

To address the concern about contributions to pensions and health plans, we compute an hourly wage measure from the 'wages and salary' portion of the QCEW, which excludes these supplements. We find that the increase in hourly wage volatility is almost identical (see appendix). To address the concern about stock options and more generally the role of high-earning individuals in the LPC, we compute an alternative hourly wage measure from the Current Population Survey (CPS). The CPS is the BLS' monthly household survey and collects a large array of labor market information, including on compensation and hours worked. A distinct advantage of the CPS for our purpose is that tips, commissions and bonuses are recorded only if they are part of regular earnings and that the publicly available data is topcoded. Stock options and, more generally, large swings in compensation of high-earning individuals therefore leave average compensation unaffected, which makes the CPS a more conservative measure for wage volatility. A disadvantage of the CPS is that hourly wages can be computed only on an annual frequency from 1973 onwards with the introduction of the CPS May supplements. Starting in 1979 then, hourly wage information is available monthly from the Outgoing Rotation Groups (ORG) files.[10] Following Lemieux (2006) and others, we construct an annual series of the average hourly wage using the May supplements for 1973-1978 and annual averages of the monthly ORG files for 1979-2006. From the total May/ORG sample, we remove all unemployed, self-employed, individuals under 16 years old, private household workers, agricultural workers and armed force personnel in order to obtain a non-farm business equivalent measure.[11] Furthermore, we multiply topcoded earnings by a constant factor of 1.3 to adjust for changes in the topcode over time and augment the average wage of hourly paid workers for 1973-1993 with a linear trend to correct for possible discontinuities arising from the 1994 CPS redesign.[12]

[10] An interviewed individual appears in the CPS for two periods of four consecutive months, separated by eight months during which the individual is left out of the survey. Before 1979, the earnings questions were asked only once a year (the May supplements). Thereafter, the earnings questions are asked each month to the individuals who are at the end of a four-month rotation (the ORGs).

[11] For 1973-78, we obtain an average of 30406 individual data points per year from the May supplements. From 1979 onwards, the combination of 12 months of ORG files into annual files yields an average of 139230 invididivual data points per year. Measurement error should therefore not be an issue. If at all, measurement error is smaller in the post-84 sample, which would lead to an understatement of the increase in wage volatility. The March supplements of the CPS provide another source of information for labor earnings. This data would have the advantage that it starts in 1963. However, the March supplements only started to collect information on total hours worked in 1976, which makes it impossible to compute hourly wages before that year. Furthermore, Lemieux (2006) argues that the earnings data from the March supplements are subject to other measurement issues not present in the CPS May/ORG files. See his paper for a detailed discussion.

[12] The appendix discusses these issues in detail. All results are robust to other topcode adjustments and corrections

The top panel of Table 2 presents the results for hourly wages from the CPS along with the annualized series for output and hourly wages from the LPC.[13] Both absolute and relative volatility of the LPC measure increase in similar proportions than in Table 1 even though the series are annualized and the pre-84 sample period only starts in 1973. The absolute volatility of the CPS measure is almost identical to the volatility of the LPC measure for the pre-84 period and also increases after 1984. This increase is, however, smaller and insignificant, suggesting that stock options and large wage changes at the top end of the distribution drive part of the increase in wage volatility in the LPC. At the same time, the volatility of the CPS wage *relative* to the volatility of output still displays an almost three-fold increase. The increase in relative wage volatility is therefore not an artifact of the increased importance of stock options, neither is it driven by the emergence of a small group of high-earning individuals with very volatile wages.

Another popular measure of hourly wages is Average Hourly Earnings (AHE) from the BLS' Current Establishment Survey (CES), which is available monthly and starts in 1964. The lower panel of Table 2 presents the results for the AHE averaged to quarterly frequency along with the corresponding series for output and LPC hourly wages. The absolute and relative volatility of the LPC wage increase in similar proportions than before, confirming the robustness of our main results for different sample periods. By contrast, the absolute volatility of the AHE is higher in the pre-84 period and then declines significantly in the post-84 period, so much that its volatility relative to the volatility of output becomes slightly smaller. This stark difference in hourly wage dynamics is not limited to the business cycle. As Abraham et al. (1998) document in earlier work, the AHE also diverges greatly from other hourly wage measures in terms of its trend. For example, whereas the LPC hourly wage increases by about 7% between 1973 and 1993, the AHE falls by about 10% over the same period. Given the frequent use of the AHE in both academic research and the business press, it is important to investigate this striking difference in results and explain why we prefer the LPC and CPS measures of hourly wages.

Conceptually, differences between the AHE measure and the LPC measure of hourly wages can come from either total compensation, total hours or both. However, the business cycle components of total hours in the CES (which underlies the AHE) and the LPC are almost identical.[14] The divergent business cycle dynamics of the AHE and the LPC wage must therefore be due to differences in total compensation. As described above, the LPC wage is based on total compensation

for the 1994 redesign.

[13] All series in Table 2 are HP filtered, with the constant set to 6.25 for annual data as recommended by Ravn and Uhlig (2002). Results are robust to alternative filters.

[14] This should not come as a big surprise since total hours in the LPC are constructed primarily from CES hours, supplemented by information from the CPS. See the appendix for details.

from the QCEW, which covers 98% of U.S. establishments and jobs. By contrast, the AHE is constructed from compensation for production and nonsupervisory workers as reported by the sample of establishments in the CES. The AHE therefore covers only a subset of the U.S. workforce that represents approximately 60% of total private-sector compensation.[15] This lack of representativeness of the AHE together with the fact that compensation in the LPC is based on a near census of establishments rather than just a sample makes for a very strong argument in favor of the LPC.

Determining the exact reasons for the divergence of the AHE remains of course an important exercise that is, however, complicated by many data issues. We pursue this investigation in a separate paper (Champagne and Kurmann, 2011) and summarize here the main findings. To assess the role played by the lack of representativeness of the AHE, we follow Abraham et al. (1998) and use the occupational information provided in the CPS to recreate an hourly wage series for production and non-supervisory workers from the May/ORG dataset.[16] Abraham et al. (1998) show that the thus simulated AHE accounts for about 60% of the above mentioned decline in the actual AHE between 1973 and 1993. Using exactly the same code as Abraham et al. (1998), we repeat their exercise and find that the simulated AHE almost exactly replicates the higher volatility of the actual AHE in the pre-84 period and generates about 35% of the decline in volatility reported in Table 3. Hence, the lack of representativeness of the AHE seems to play a substantial role for its divergence, both in terms of trend and volatility, from other wage measures.

The second issue with the AHE measure examined in Champagne and Kurmann (2011) concerns sampling problems. In particular, the CES sample underwent a substantial expansion from about 160,000 to about 400,000 establishments between 1980 and 2006. A large part of that expansion occurred for young and small establishments in service industries, which were severely underrepresented in the early 1980s (e.g. Plewes, 1982).[17] The resulting improvement in the sample properties of the CES is likely to have led to spurious changes in the AHE measure, both because measurement errors for wages in service industries decreased substantially and because small and young firms in service industries hire on average less skilled workers for which hourly wages have become less volatile (see next section).[18] None of these sampling problems are present in either the LPC or

[15] According to Abraham et al. (1998), production and non-supervisory workers account for about 80% of private-sector employment. But because both their hourly wage and hours worked are lower than the economy-wide average, compensation of production and non-supervisory workers only represents about 60% of private-sector compensation.

[16] The definition of compensation in the AHE and the CPS is very similar. Both of them record commission and bonuses only if earned and paid during the same period. Compensation in the CPS also includes tips whereas AHE does not. As Abraham et al. (1998) report, however, tips to account for a mere 0.3% of total compensation in 1993.

[17] Note that employment numbers in the CES are benchmarked to the UI records, which are the source of the QCEW. Earnings and hours are, however, not benchmarked and therefore do not undergo a regular bias correction.

[18] Another sampling problem with the CES is that it captures the birth and death of establishments only with a

the CPS. Together with the lack of representativeness of the AHE, these sampling problems lead us to conclude that the LPC and CPS measures of hourly wages should be unambiguously preferred over the AHE.

2.4 Other changes in labor market dynamics

A number of recent papers document that the Great Moderation has been accompanied by substantial changes in the business cycle dynamics of aggregate hours and labor productivity. To put our results into perspective, we report several of these changes in Table 3 and compare them to changes in hourly wage dynamics. The first four rows of the table show relative volatilities for the pre-84 and post-84 period. The relative volatility of both aggregate hours and labor productivity increases, a result first uncovered by Gali and Gambetti (2009).[19] Compared to the more than three-fold increase in the relative volatility of real hourly wages, these changes are, however, modest. Interestingly, the relative volatility of *nominal* hourly wages also increases substantially but not quite as much as the relative volatility of real hourly wages.

The last four rows of Table 3 report different correlation coefficients. The correlation of labor productivity with both output and hours declines substantially during the Great Moderation, a phenomenon documented in Stiroh (2009) and Gali and Gambetti (2009).[20] This decline in cyclicality also applies to real hourly wages, although to a somewhat lesser extent. Finally, the correlation of nominal hourly wages with prices turns from highly positive to almost zero. To our knowledge, this result is new and provides an interesting perspective for the increase in volatility of real hourly wages. Since real hourly wages equal nominal hourly wages divided by the price level, the increase in the volatility of real hourly wages can be decomposed into changes in the volatility of nominal hourly wages, the volatility of the price level and the negative of the correlation between the two variables.[21] The (absolute) volatility of nominal hourly wages has remained roughly constant and the volatility of the price level has fallen substantially. Hence, the increase in the volatility of real hourly wages is to a large part accounted for by the drop in the correlation between nominal hourly wages and the price level.

time-lag. CES employment numbers are adjusted for the resulting sampling bias but earnings and hours are not. This is likely to affect the cyclical properties of the AHE because new and disappearing establishments pay different hourly wages than continuing establishments. The expansion of the CES with mostly young and small firms may also have changed the birth/death bias and thus the volatility of the AHE.

[19] Labor productivity is computed as output divided by total hours from the LPC.

[20] In related work, Barnichon (2010) shows that the correlation of labor productivity with unemployment switched from mildly negative to significantly positive for the same sample periods.

[21] The variance of the business cycle component of log real hourly wages is approximately $var(w) \approx var(w^{nom}) + var(p) - 2corr(w^{nom}, p) \times \sqrt{var(w^{nom})var(p)}$

3 A closer look at disaggregated data

The documented increase in the relative volatility of hourly wages during the Great Moderation raises two important questions. First, did this increase occur for all groups or types of workers? Second, what is the role played by changes in workforce composition? The answer to these questions provides valuable clues in the search for possible explanations. If, for example, the relative volatility of hourly wages increases in similar proportions for many groups of workers, then this directs us towards changes in the economic environment that affect different labor markets alike. If, to the contrary, the relative volatility of hourly wages remains approximately constant for most groups of workers, then we need to focus on other explanations such as the role played by changes in workforce composition towards jobs with more volatile wages.

3.1 Wage volatility across different decompositions

The CPS May/ORG dataset provides a wealth of information for all interviewed individuals that we can use to decompose the workforce into groups with different characteristics. Following Krusell et al. (2000) and many others, we choose education as one of the characteristics, with a 'skilled worker' being someone with a college degree (bachelor) or more, and an 'unskilled worker' being someone with less than a college degree. On top of education, we add in rotating order the following distinctions: gender, age, compensation status (hourly paid or salaried), and industry affiliation. This yields four different decompositions: (i) gender / education; (ii) age / education; (iii) compensation status / education; and (iv) industry affiliation / education.[22] For each of the decompositions, we compute an average hourly wage series and follow the same procedure as above: filter the series to extract the business cycle component; split the sample into a pre-1984 and a post-1984 period; and compute the volatility of the hourly wage series both in absolute terms and relative to the volatility of aggregate output.

In light of the prominent role played by deunionization in the model of Section 4, it would be interesting to do a decomposition along union membership as well. Unfortunately, the CPS ORG files do not contain information on union membership before 1983; for 1982 there is no union information at all; and for 1979 to 1981, the number of individuals in the CPS May with both wage and union information is only about one quarter of the regular sample and not representative of the U.S. workforce.[23] Since the years 1979-1983 are very important for the determination of wage

[22] In the appendix, we report an additional decomposition with respect to gender and occupation which yields similar results than the decomposition for gender and education. It would also be interesting to decompose the CPS sample further along more than just two characteristics. Cell size becomes quickly an issue, however.

[23] More specifically, the CPS May supplements asked the union question to all workers in 1979 and 1980 but only

volatility in the pre-84 period, these data problems make it impossible to compute reliable results for a union / non-union decomposition.

Gender / education decomposition. The first panel of Table 4 reports the results for the gender / education decomposition. The absolute volatility of hourly wages increases strongly for skilled males, stays approximately constant for unskilled males and females, and drops for skilled females. Relative to the volatility of output, the volatility of hourly wages increases substantially across all groups. Most notable is the more than six-fold increase for skilled male workers.

Age / education decomposition. The second panel of Table 4 displays the results for the age / education decomposition. Following Gomme et al. (2004), and Jaimovich and Siu (2008), we create three age groups: 16-29 year olds ('young workers'); 30-59 year olds ('middle-age workers'); and 60-70 year olds ('old workers'). The absolute volatility of hourly wages increases for both young and middle-aged skilled workers, remains approximately constant for young and middle-aged unskilled workers, and declines for old workers. As in the previous decomposition, the relative volatility of hourly wages increases substantially for all groups, especially for the young and middle-aged skilled workers.

Compensation status / education decomposition. The third panel of Table 4 shows the results for the employment status / skill decomposition, where employment status is defined by whether a worker is salaried or paid on an hourly basis. The absolute volatility of hourly wages remains constant or falls slightly for all but the salaried skilled group, for which the volatility doubles. The relative volatility of hourly wages increases again markedly for all worker groups and is most pronounced for the salaried skilled group.

Industry / education decomposition. Table 5 reports the results for the industry affiliation / education decomposition. We choose a relatively detailed decomposition into ten private sector industries.[24] The absolute volatility of hourly wages increases for unskilled workers employed in manufacturing and other services and for skilled workers in wholesale trade and finance, insurance and real estate (FIRE). For all other groups, the absolute volatility of the hourly wage decreases. Except for workers employed in communications, this decrease in volatility is modest and thus, the increase in the relative volatility of wages remains pervasive. The marked decrease in wage volatility for the communications industry is noteworthy because this is an industry where unionization remains high relative to other private-sector industries (over 20% compared to less than 10% on

to one quarter of the sample in 1981. The earnings question, in turn, was asked to half the sample in 1979 and one quarter of the sample in each 1980 and 1981. In addition, the answer for the earnings question is missing for a substantial portion of the queried sample. See Hirsch and Schumacher (2004) for details.

[24]See appendix for industry classification. Note that for this decomposition, the sample stops in 2002 because the industry reclassification for 2003 (from SIC to NAICS) makes matching of some 3-digit industries difficult.

average). Interestingly, the only other sector where wage volatility declines even more than in communications is the public sector – a sector where unionization increased to almost 40% during the post-84 period. While our focus is on wage dynamics in the private sector, we think that this result is indicative of the importance of unionization for wage setting.

We take away two stylized facts from these decompositions. First, there is substantial heterogeneity in how the absolute volatility of hourly wages of different worker groups changes over time. The largest increases in volatility occur for skilled workers that are either male, young or middle-aged or salaried. Many other groups, especially in the industry / education decomposition, see the volatility of their hourly wage decrease. Second, the decrease in wage volatility for these latter groups is generally modest relative to the decrease in the volatility of output. Hence, the volatility of hourly wages *relative* to the volatility of output increases substantially for almost all worker groups. This phenomenon is what we call in the introduction 'The Great Increase in Relative Volatility of Real Wages'.

3.2 Volatility accounting

While the increase in relative wage volatility is pervasive across worker groups, it might still be the case that a large part of the increase in the relative volatility of average hourly wages is driven by secular changes in workforce composition towards jobs where hourly wages are typically more volatile. To assess this question, we develop a volatility accounting method that allows us to quantify how much of the increase in the relative volatility of average hourly wages is due to changes in workforce composition and how much is due to increases in the relative volatility of hourly wages across the different worker groups.

By definition, the average hourly wage w_t equals the sum of hourly wages $w_{i,t}$ across worker groups i of some decomposition (e.g. gender / education), weighted by the respective hours shares $h_{i,t} = H_{i,t}/H_t$; i.e. $w_t = \sum_i w_{i,t}h_{i,t}$. Next, let $x_{i,t} \equiv w_{i,t}h_{i,t}$ be the 'wage component of group i' and express the growth rate of the aggregate wage as

$$\Delta \log w_t \approx \frac{w_t - w_{t-1}}{w_{t-1}} = \sum_i \frac{x_{i,t-1}}{w_{t-1}} \frac{x_{i,t} - x_{i,t-1}}{x_{i,t-1}} \approx \sum_i s_{i,t-1} \Delta \log x_{i,t} = \sum_i s_{i,t-1}(\Delta \log w_{i,t} + \Delta \log h_{i,t}),$$

where $s_{i,t-1} = x_{i,t-1}/w_{t-1} = (w_{i,t-1}H_{i.t-1}) / (w_{t-1}H_{t-1})$ denotes the 'wage share' of worker group i; i.e. the weight with which group i's hourly wage growth (or its hours share growth) affects average hourly wage growth. Since the wage shares of the different worker groups evolve slowly over time, we assume that $s_{i,t-1}$ is well approximated by its average for the subsample under consideration (say, the pre-84 period); i.e. $s_{i,t-1} \approx \bar{s}_i$. This allows us to write the variance of average hourly wage growth relative to the variance of output as

$$\frac{\sigma_w^2}{\sigma_y^2} \approx \sum_i \bar{s}_i^2 \left[\frac{\sigma_{w_i}^2}{\sigma_y^2} + \frac{\sigma_{h_i}^2}{\sigma_y^2} + 2\frac{\sigma_{w_i,h_i}}{\sigma_y^2} \right] + 2\sum_{i \neq j} \bar{s}_i \bar{s}_j \left[\frac{\sigma_{w_i,w_j}}{\sigma_y^2} + \frac{\sigma_{h_i,h_j}}{\sigma_y^2} + \frac{\sigma_{w_i,h_j}}{\sigma_y^2} \right],$$

where $\sigma_w^2 \equiv var(\Delta \log w_t)$, $\sigma_{w_i,h_j} \equiv cov(\Delta \log w_{i,t}, \Delta \log h_{j,t})$ and so forth for the other terms. We check this approximation for each of the decompositions and find the induced error to be negligible. The decomposition makes clear that by acting as weights on the different relative variance and covariance terms, the wage shares of the different worker groups play an important role for the determination of the relative variance of average hourly wage growth. Given this decomposition, the change in the relative variance of average hourly wage growth from one subsample a to another subsample b (i.e. from the pre-84 period to the post-84 period) equals[25]

$$\begin{aligned} \frac{\sigma_w^2(b)}{\sigma_y^2(b)} - \frac{\sigma_w^2(a)}{\sigma_y^2(a)} \approx\ & \sum_i \bar{s}_i^2(b) \left[\frac{\sigma_{w_i}^2(b)}{\sigma_y^2(b)} + \frac{\sigma_{h_i}^2(b)}{\sigma_y^2(b)} + 2\frac{\sigma_{w_i,h_i}(b)}{\sigma_y^2(b)} \right] - \sum_i \bar{s}_i^2(a) \left[\frac{\sigma_{w_i}^2(a)}{\sigma_y^2(a)} + \frac{\sigma_{h_i}^2(a)}{\sigma_y^2(a)} + 2\frac{\sigma_{w_i,h_i}(a)}{\sigma_y^2(a)} \right] \\ & +2\sum_{i \neq j} \bar{s}_i(b)\bar{s}_j(b) \left[\frac{\sigma_{w_i,w_j}(b)}{\sigma_y^2(b)} + \frac{\sigma_{h_i,h_j}(b)}{\sigma_y^2(b)} + \frac{\sigma_{w_i,h_j}(b)}{\sigma_y^2(b)} \right] \\ & -2\sum_{i \neq j} \bar{s}_i(a)\bar{s}_j(a) \left[\frac{\sigma_{w_i,w_j}(a)}{\sigma_y^2(a)} + \frac{\sigma_{h_i,h_j}(a)}{\sigma_y^2(a)} + \frac{\sigma_{w_i,h_j}(a)}{\sigma_y^2(a)} \right]. \end{aligned} \tag{1}$$

Hence, the change in the relative variance of the average hourly wage growth can come from four different sources: (i) changes in average wage shares (i.e. the effect of compositional changes discussed above); (ii) changes in the relative variance of hourly wage growth of the different groups; (iii) changes in the relative variance of the hours share growth of the different groups; and (iv) changes in different correlation coefficients (since changes in the various covariance terms in (1) are themselves a function of changes in the respective correlation and relative variance terms; e.g. for the covariance between two variables z_1 and z_2, we have $\sigma_{z_1,z_2} = corr_{z_1,z_2}\sigma_{z_1}\sigma_{z_2}$). Our volatility accounting exercise consists of decomposing the total change in the relative variance of average hourly wage growth into these four sources.

Since the different wage shares, variances and correlations on the right-hand side of (1) enter in multiplicative combinations, we need to take a stand on the 'base period' with which to weigh each of the four sources, keeping everything else constant. To see this, we reintroduce the definition of

[25] While our focus is on accounting for the increase in the *relative* volatility of hourly wages, the same exercise could be performed for the *absolute* volatility. We know from Section 2, however, that the increase in absolute volatility of hourly wages was at most modest (according to the CPS data, which we use here again). Decomposing this small increase in volatility into its different sources would not be very informative.

the wage component $x_{i,t} \equiv w_{i,t}h_{i,t}$ and rewrite (1) in more compact form as

$$\frac{\sigma_w^2(b)}{\sigma_y^2(b)} - \frac{\sigma_w^2(a)}{\sigma_y^2(a)} \approx \sum_i \sum_j \bar{s}_i(b)\bar{s}_j(b)\frac{\sigma_{x_i,x_j}(b)}{\sigma_y^2(b)} - \sum_i \sum_j \bar{s}_i(a)\bar{s}_j(a)\frac{\sigma_{x_i,x_j}(a)}{\sigma_y^2(a)}. \tag{2}$$

By adding and subtracting elements, we can expand this equation in two different ways

$$\begin{aligned} \frac{\sigma_w^2(b)}{\sigma_y^2(b)} - \frac{\sigma_w^2(a)}{\sigma_y^2(a)} &\approx \sum_i \sum_j \left[\bar{s}_i(b)\bar{s}_j(b) - \bar{s}_i(a)\bar{s}_j(a)\right] \frac{\sigma_{x_i,x_j}(a)}{\sigma_y^2(a)} \\ &\quad + \sum_i \sum_j \bar{s}_i(b)\bar{s}_j(b) \left[\frac{\sigma_{x_i,x_j}(b)}{\sigma_y^2(b)} - \frac{\sigma_{x_i,x_j}(a)}{\sigma_y^2(a)}\right] \\ &\approx \sum_i \sum_j \left[\bar{s}_i(b)\bar{s}_j(b) - \bar{s}_i(a)\bar{s}_j(a)\right] \frac{\sigma_{x_i,x_j}(b)}{\sigma_y^2(b)} \\ &\quad + \sum_i \sum_j \bar{s}_i(a)\bar{s}_j(a) \left[\frac{\sigma_{x_i,x_j}(b)}{\sigma_y^2(b)} - \frac{\sigma_{x_i,x_j}(a)}{\sigma_y^2(a)}\right]. \end{aligned} \tag{3}$$

The first expansion decomposes the change in the relative variance of average hourly wage growth into changes in wage shares *weighted by the covariances of the wage components for the first subsample* and changes in covariances of the wage components *weighted by the wage shares for the second subsample*. The second expansion decomposes the relative variance of the average hourly wage growth into changes in wage shares *weighted by the covariances of the wage components for the second subsample* and changes in covariances of the wage components *weighted by the wage shares for the first subsample*. Since there is no economic justification to prefer one 'base period' over the other, we take the average over the two expansion and obtain[26]

$$\begin{aligned} \frac{\sigma_w^2(b)}{\sigma_y^2(b)} - \frac{\sigma_w^2(a)}{\sigma_y^2(a)} &\approx \sum_i \sum_j \left[\bar{s}_i(b)\bar{s}_j(b) - \bar{s}_i(a)\bar{s}_j(a)\right] \left[\frac{\frac{\sigma_{x_i,x_j}(b)}{\sigma_y^2(b)} + \frac{\sigma_{x_i,x_j}(a)}{\sigma_y^2(a)}}{2}\right] \\ &\quad + \sum_i \sum_j \left[\frac{\bar{s}_i(b)\bar{s}_j(b) + \bar{s}_i(a)\bar{s}_j(a)}{2}\right] \left[\frac{\sigma_{x_i,x_j}(b)}{\sigma_y^2(b)} - \frac{\sigma_{x_i,x_j}(a)}{\sigma_y^2(a)}\right]. \end{aligned} \tag{4}$$

While this specification of base period for the weighting is arbitrary, we note that none of the substantive results of our volatility accounting exercise would change if we instead used the first or the second expansion in (3). The first line on the right-hand side of (4) captures the effect of changes in workforce composition on the relative variance of average hourly wages. The second line captures the effect of changes in the relative variances and correlations of the different $x_{i,t}$ terms. Using the

[26] This problem of choosing a 'base period' is conceptually similar to the problem faced in national accounting when computing *real* macro aggregates (e.g. real GDP). Our approach to use an average as the 'base period' for the weights resembles the chain-type method used in the NIPAs.

fact that $\Delta \log x_{i,t} = \Delta \log w_{i,t} + \Delta \log h_{i,t}$ and applying the same averaging of 'base periods', we can expand this second part further into a weighted sum of changes in the relative variances of hourly wage growth of the different groups, changes in the relative variances of hours share growth of the different groups and changes in the various correlation terms. Since this expansion involves tedious algebra, we delegate the details to the appendix. The end result is an additive expression that allows us to decompose the relative variance of average hourly wage growth into the four sources described in (i) through (iv).

Table 6 reports the results of this volatility accounting exercise for each of the four decomposition analyzed above.[27] The second line shows that changes in wage shares play only a modest role, contributing at most 13% to the increase in the relative variance of average hourly wages. Changes in the relative variance of hourly wages of the different worker groups, by contrast, explain 70% or more of the increase in the relative variance of average hourly wages. The rest is accounted for by changes in the relative variance of hours shares and changes in the different correlation coefficients.[28] This shows that the widespread increase in the relative volatility of hourly wages across different worker groups is the main source of the increase in the relative volatility of average hourly wages. The absence of sizable effects from structural changes in workforce composition directs the search for possible explanations towards changes in the economic environment that have similar effects on wage setting in different labor markets.[29] At the same time, some worker groups experience a larger increase in relative wage volatility than others, which suggests that these structural changes do not occur to the same extent for everyone.

[27] All results pertain to HP-filtered data. This introduces an additional approximation error since the decomposition in (1) is based on growth rates. We find that this error is minimal for all decompositions considered.

[28] Interestingly, changes in the relative variances of hours shares of the different worker groups have a small negative effect. In the appendix, we show that the volatility of hours shares declined markedly for many worker groups in our decomposition, which explains this result.

[29] Since the CPS data does not allow us to follow individual workers over time, we cannot rule out that compositional effects play a role *within* worker groups. Starting with Gottschalk and Moffitt (1994), however, different papers using panel data on individual workers show that labor income has become considerably more volatile (in absolute terms) for some individuals (mostly the ones with high incomes) and has remained roughly constant for the rest. See Dynan et al. (2008) and Jensen and Shore (2008) for recent contributations and an extensive review. Since the volatility of output fell by more than 50% during the same time period, this means that the *relative* volatility of labor income must have increased substantially on an individual level as well. This provides additional support for our conclusion that compositional effects play only a minor role for the increase in relative volatility of average hourly wages.

4 Wage volatility in general equilibrium

In this section, we develop a DSGE model to quantitatively assess two possible explanations for the increase in relative wage volatility. First, we explore the effects of the 'good luck hypothesis', i.e. a decrease in the importance of exogenous shocks, that many studies credit as the main driver of the Great Moderation.[30] Second, we consider two structural changes in the labor market: deunionization and increased incidence of performance-pay.

Our focus on deunionization and increased incidence of performance-pay is motivated by a combination of empirical observations. First, the U.S. labor market experienced a marked decline in unionization and a shift towards performance-pay contracts over the past decades. Figure 2 illustrates these developments. The left panel plots the proportion of non-union workers in the non-farm business sector computed from data in Hirsch et al. (2001) and Hirsch and Macpherson (2010).[31] The right panel plots a measure of the incidence of performance-pay from Lemieux et al. (2009a), defined as the proportion of male household heads in the PSID whose compensation in an employment relationship includes a variable pay component (bonus, commission, or piece-rate).[32] As the plots show, both deunionization and the shift towards performance-pay accelerate in the early 1980s and then continue to rise, although at a lower pace, during the 1990s.[33] This acceleration coincides approximately with the increase in hourly wage volatility documented above. In addition, the above studies document that deunionization and increased incidence of performance-pay occurs (to varying degrees) for most of the different worker groups considered, which is consistent with our conclusion above that the increase in relative wage volatility is widespread and not concentrated in a small segment of the U.S. labor market.

[30] See for example Stock and Watson (2002) or Sims and Zha (2006).

[31] Hirsch et al. (2001) compute union density for the entire U.S. economy between 1964-2000 by combining information from the BLS publication Directory of National Unions and Employee Associations for 1964-72 and from the CPS May/ORG from 1973 onward. They also adjust all CPS May/ORG numbers before 1977 to reflect a change in the union question in the CPS. To obtain union density for the non-farm business sector, we take the non-farm business union density series reported in Hirsch and Macpherson (2010) for the years 1973-2006 and assume that the ratio of union-density in the non-farm business sector to union density in the entire economy between 1964 and 1976 is the same than in 1977. This probably implies too conservative of an estimate of union density before 1977. For 1982, where no union information is available, we linearly interpolated non-union density from information in 1981 and 1983.

[32] Unfortunately, the CPS data does not have information on performance-pay.

[33] As we argue in the calibration section below, Lemieux et al.'s (2009a) PSID measure of performance-pay is likely to underestimate the true increase performance-pay contracts. Other studies discussing the causes and effects of increases in performance-pay contracts in the U.S. are Mitchell et al. (1990), Prendergast (1999) and Cunat and Guadalupe (2005, 2008).

Second, based on the same PSID dataset, Lemieux et al. (2009b) show that wages of non-union workers with performance-pay contracts are most responsive to local labor market shocks and least responsive for union workers without performance-pay. Exactly the opposite is the case for hours worked, suggesting that wages play an allocative role over the business cycle. Furthermore, Lemieux et al. (2009b) report that workers with performance-pay are on average more skilled and more likely to be in salaried positions. Our analysis of the CPS data above reveals that wage volatility increased most for exactly these types of workers.

Taken together, these empirical observations suggest that deunionization and the shift towards performance-pay result in greater wage flexibility, making wages more and hours (and output) less responsive to business cycle shocks. We use our DSGE model to assess this conjecture in general equilibrium and quantify to what extent these labor market developments contribute to the observed changes in wage and output volatility during the Great Moderation.

4.1 Model

The model contains many elements that are standard in the DSGE literature. In particular, as in New Keynesian models with nominal wage rigidities, we assume that each worker supplies a differentiated labor service and sets wages according to a given contract rule. Based on these wages, firms then choose the optimal combination of labor services to minimize labor costs. The key novelty here is that we extend this wage setting idea so as to distinguish between union and non-union workers who are hired either on performance-pay or non-performance-pay contracts. This allows us to simulate the effects of deunionization and increased incidence of performance-pay.

The economy is populated by three types of agents: a continuum of infinitely-lived workers; a continuum of infinitely-lived firms; and a government that determines monetary and fiscal policy. Workers discount time at rate β and have preferences over consumption and leisure. Total expected lifetime utility of worker i is

$$\mathbf{E}_0 \sum_{t=0}^{\infty} \beta^t Z_{t-1} \left[\log C_t - \frac{N_t(i)^{1+\phi}}{1+\phi} \right], \tag{5}$$

and the per period budget constraint is

$$C_t + K_{t+1} - (1-\delta)K_t + \frac{B_{t+1}}{R_t^n P_t} + T_t \leq \frac{W_t(i)N_t(i)}{P_t} + R_t^K K_t + \frac{B_t}{P_t} + D_t + \Gamma_t(i). \tag{6}$$

$\mathbf{E}_0$ denotes the expectations operator; Z_{t-1} an exogenous preference shock common to all workers; C_t a composite consumption good; $N_t(i)$ hours worked; $K_{t+1} - (1-\delta)K_t$ investment in physical capital; B_t nominal bond holdings; T_t lump-sum taxes; D_t dividends from a perfectly diversified portfolio of claims to firms; $\digamma_t(h)$ the net return from a state-contingent insurance mechanism; $W_t(i)$ the nominal wage rate; R_t^K the real net rental rate of capital; R_t^n the gross nominal bond return; and P_t the aggregate price level. Labor income $W_t(i)N_t(i)$ is worker-specific due to the labor market frictions described below. As in Erceg et al. (2000), the net return $\digamma_t(i)$ is such that workers remain identical with respect to their consumption and savings decisions.[34] This is why our notation for C_t, K_t and B_t is not worker-specific.

Each worker supplies a differentiated labor service and either belongs to a union or not. Firms produce with a labor composite N_t that is made up of union labor N_t^u and non-union labor N_t^{nu} according to the aggregator

$$N_t = \left[s^u (N_t^u)^{\frac{\mu-1}{\mu}} + s^{nu} (N_t^{nu})^{\frac{\mu-1}{\mu}} \right]^{\frac{\mu}{\mu-1}}, \tag{7}$$

where s^u and $s^{nu} \equiv 1 - s^u$ are fixed weights that pin down the average wage shares of the union sector and the non-union sector; and $\mu > 1$ is the elasticity of substitution determining the extent to which firms can switch between union and non-union labor over the business cycle. Union and non-union labor are themselves a Dixit-Stiglitz aggregate of the differentiated labor services of union and non-union workers, respectively

$$N_t^l = \left[\int_0^1 N_t^l(i)^{\frac{\mu^l-1}{\mu^l}} di \right]^{\frac{\mu^l}{\mu^l-1}} \quad \text{for } l \in \{u, nu\}. \tag{8}$$

The elasticities $\mu^u > 1$ and $\mu^{nu} > 1$ determine the extent to which union workers and non-union workers, respectively, are substitutable among each other. Given (7) and (8), the firms' optimal labor demand for a union worker charging $W_t^u(i)$ and for a non-union worker charging $W_t^{nu}(i)$ is

$$N_t^l(i) = \left(\frac{W_t^l(i)}{W_t^l} \right)^{-\mu^l} \times \left(\frac{1}{s^l} \frac{W_t^l}{W_t} \right)^{-\mu} N_t \;\; \text{for } l \in \{u, nu\}, \tag{9}$$

where W_t^u and W_t^{nu} denote the aggregate union and non-union wage; and W_t is the aggregate index of the labor composite N_t that firms use to produce.

For wage setting, we assume that a fraction p^u of union workers and a fraction p^{nu} of non-union workers receive performance-pay. The defining feature of a performance-pay contract is that part or all of the compensation is linked to observed output by the worker (see Lemieux et al., 2009b for an illustrative model and a review of the literature).[35] In the context of our model, we take

[34] This type of insurance mechanism is common in DSGE models with heterogenous labor market outcomes and can alternatively be achieved through a household structure as in Andolfatto (1996).

[35] In Lemieux et al. (2009b), performance-pay contracts arise endogenously if the costs of overcoming informational frictions are sufficiently small. Otherwise, firms and workers find it optimal to not measure performance and enter a fixed-wage contract.

this feature to mean that the nominal wage $W_t^{l,p}(i)$ of a performance-pay worker adjusts with time t information so as to satisfy the labor supply condition

$$\frac{W_t^{l,p}(i)}{P_t} = \frac{\mu^l}{\mu^l - 1} N_t^{l,p}(i)^{\phi} C_t \quad \text{for } l \in \{u, nu\}, \tag{10}$$

where $N_t^{u,p}(i)$ and $N_t^{nu,p}(i)$ are determined by the firms' respective labor demands in (9). Otherwise, either the wage or hours worked would need to adjust for the contract to remain incentive compatible.[36]

For the remaining workers without performance-pay, we assume that nominal wages are set in advance of time t information according to a variant of Calvo (1983). In the union sector, the fraction of non-performance-pay workers (or equivalently, the fraction of unions) that get to reoptimize their nominal wage for next period is $1 - \xi^u$. In the non-union sector, the equivalent fraction is $1 - \xi^{nu}$. For all other non-performance pay workers (a fraction ξ^u in the union sector and a fraction ξ^{nu} in the non-union sector), wages are indexed to the steady state growth rate of consumption γ and partially to realized gross inflation Π_{t-1}; i.e. their nominal wage adjusts according to $W_t^{l,np}(i) = \gamma \Pi_{t-1}^{\omega} W_{t-1}^{l,np}(i)$ with ω denoting the inflation indexing factor. The optimal wage contract of a non-performance-pay worker who gets to reoptimize for time t therefore solves

$$\max_{W_t^{l,np}(i)} E_{t-1} \sum_{j=0}^{\infty} (\beta \xi^l)^j \left[\frac{1}{C_{t+j}} \frac{W_t^{l,np}(i) X_{t,t+j}}{P_{t+j}} N_{t+j}^{l,np}(i) - \frac{N_{t+j}^{l,np}(i)^{1+\phi}}{1+\phi} \right] \quad \text{for } l \in \{u, nu\} \tag{11}$$

subject to labor demand (9); with $X_{t,t+j} \equiv \prod_{s=1}^{j} \gamma \Pi_{t+s-1}^{\omega}$ for $j \geq 1$ and $X_{t,t+j} = 1$ for $j = 0$.

Several comments are in order about this formalization of wage setting. First, as in existing New Keynesian DSGE models, we abstract from the deeper frictions that give rise to staggered wage reoptimization in some employment relationships but not in others. Likewise, our performance-pay contract is not derived from an explicit principal-agent or hold-up problem as in Lemieux et al. (2009b); and we do not explicitly model the forces that lead to unionization. While very interesting, such a richer environment would exceed the objective of our model, which is to quantify the general equilibrium effects of changes in unionization and the incidence of performance-pay, everything else constant. Second, firms in our model have the right-to-manage; i.e. they can freely decide on labor input given a set wage. The right-to-manage assumption is consistent with most U.S. labor market

[36] Note that the wage of a union worker with performance-pay typically differs from the wage of a non-union worker with performance-pay. This is because the optimal markup of the wage over the marginal rate of substitution that a union worker commands (i.e. $\mu^u/(\mu^u - 1)$) is different from the optimal markup of a non-union worker (i.e. $\mu^{nu}/(\mu^{nu} - 1)$); and because in equilibrium, the average union wage W_t^u and therefore labor demand evolves differently from the average non-union wage W_t^{nu}.

contracts (see Malcomson, 1999) and the empirical results in Lemieux et al. (2009b) suggest that firms indeed adjust hours most for workers with the least flexible wages. Third and related, workers without performance-pay contracts in our model are typically not on their labor supply curve. Given the markup they command in the labor market, their wage remains, however, above the marginal rate of substitution; i.e. the wage more than compensates for the disutility from working.

The rest of the model is standard. We thus keep the exposition to a minimum and refer the reader to the appendix for a full description. Given labor and investment income, workers in each period consume and invest in either physical capital or nominal bonds to maximize (5) subject to (6). The resulting optimality conditions yield the log-linearized Euler equations

$$c_t = E_t c_{t+1} - \frac{\beta r^k}{\gamma} E_t r^k_{t+1} - \Delta z_t \tag{12}$$

and

$$c_t = E_t c_{t+1} - (r^n_t - E_t \pi_{t+1}) - \Delta z_t, \tag{13}$$

where, from here on, lower-case variables denote percent-deviations from appropriately normalized steady states. On the production side, we assume that each firm uses Cobb-Douglas technology

$$y_t = a_t + \alpha k_t + (1-\alpha) n_t, \tag{14}$$

where a_t is an exogenous technology shock common to all firms. The different firms' goods are imperfectly substitutable and are combined by a wholesale firm into the final composite good according to the Kimball (1995) aggregator.[37] Price setting follows Calvo (1983), with each firm facing a constant probability of reoptimizing its price in a given period. These frictions imply the well-known New Keynesian Phillips curve (NKPC)

$$\pi_t = \beta E_t \pi_{t+1} + \kappa mc_t, \tag{15}$$

where mc_t denotes real marginal cost; and the slope coefficient κ is a nonlinear function of price setting and demand parameters (see, for example, Eichenbaum and Fischer, 2007). Given demand, each firm minimizes costs by choosing the labor composite n_t and capital k_t. The resulting log-linearized aggregate labor and capital demand conditions are

$$w_t - p_t = mc_t + y_t - n_t, \tag{16}$$

and

$$r^k_t = mc_t + y_t - k_t. \tag{17}$$

[37] Kimball's (1995) aggregator is a generalization of the Dixit-Stiglitz aggregator and provides flexibility in mapping micro data on price adjustment to aggregate inflation dynamics. See, for example, Eichenbaum and Fischer (2007).

The government, finally, conducts monetary policy according to the following interest rate rule

$$r_t^n = \rho r_{t-1}^n + (1-\rho)[\theta_\pi \pi_t + \theta_y (y_t - y_{t-1})]; \tag{18}$$

and limits fiscal policy to a constant spending rule that is fully financed by lump-sum taxes.

4.2 Calibration

The model is calibrated to quarterly U.S. data. The structural model parameters are partitioned into two groups. The first group contains the parameters not directly related to wage setting. These parameters are kept unchanged in the different simulations and are calibrated as shown in Table 7. The values of α, β, γ, δ and ϕ are standard. The steady-state government spending-output ratio of 0.15 implies an average consumption-output ratio of 0.63 in line with the data. The NKPC slope coefficient κ lies in the range of estimates reported by different empirical studies on the NKPC. The monetary policy parameters, finally, are also in line with estimates found in the literature (e.g. Smets and Wouters, 2007). The particular values are chosen such that, conditional on the calibration of all other parameters and shock processes below, the model matches the pre-84 volatility of H-P filtered output and comes close to the correlation coefficients of labor productivity with output and hours in Table 3.

The second group of model parameters is directly related to wage setting. The wage share of union workers and the incidence of performance-pay are calibrated separately for the pre-84 and the post-84 period. We use this information in the simulations to quantify the effects of deunionization and increased incidence of performance-pay. The other parameters are calibrated over the entire sample and are kept unchanged in the simulations. Table 8 reports the different values. To calibrate the wage share of unionized workers, which pins down the parameter s^u in (7), we decompose it into three different parts: $W^u N^u/(WN) = W^u/W \times H^u/H \times E^u/E$. The ratio W^u/W denotes the hourly wage of union workers relative to average hourly wages and is related to the union wage premium W^u/W^{nu} by $W/W^u = 1 + 1/(W^u/W^{nu})$. Using annual estimates in Hirsch and Macpherson (2010, Table 2a), we obtain an average union premium in the private-sector of 1.23 for the pre-84 period and 1.24 for the post-84 period.[38] This remarkable constancy leads us to keep the union wage premium average sample value of $W^u/W^{nu} = 1.235$ over the entire sample. The ratio H^u/H denotes average hours per union worker relative to average hours over all workers. From our CPS May/ORG data, we find that this ratio averages to approximately one for both the pre-84 and

[38] Hirsch and Macpherson (2010) compute their estimates from the CPS May/ORG dataset after correcting for imputation bias in union wages and systematic differences in characteristics of union and non-union workers. See Hirsch and Schumacher (2004) for details.

the post-84 period. We therefore set $H^u/H = 1$ for the entire sample. Finally, E^u/E denotes the the proportion of union workers in the workforce (i.e. the union density). Using the union density numbers reported in Figure 2, we find that the average union density is 0.25 for the pre-84 period and 0.11 for the post-84 period.[39] Combining these numbers together yields the union wage shares in Table 6. As the calculations show, the fall in the wage share of union workers over time is entirely driven by the fall in union density.

The calibration of p^u and p^{nu} is more challenging as we do not have a direct measure of performance-pay in the CPS. Lemieux et al.'s (2009a) measure that we plot in Figure 2 can be used as a starting point. However, their sample only covers 1976 to 1998 and their measure explicitly excludes any performance-pay related to overtime work. Since performance-pay was presumably even less common before 1976 and overtime work increased substantially during the 1980s and 1990s (e.g. Kuhn and Lozano, 2008), Lemieux et al.'s (2009a) measure is likely to substantially underestimate the true increase in the incidence of performance-pay. We therefore use the available information in Lemieux et al. (2009a) and extrapolate to what we believe are reasonable values.[40]

The remaining parameters in Table 8 are calibrated as follows. The fraction of non-reoptimizing union workers ξ^u is set such that the average contract duration for union workers is 12 quarters, as reported in Rich and Tracy (2004). According to their estimates, this average remained surprisingly constant over their entire sample under consideration (1970-1995). For non-union workers, the corresponding fraction ξ^{nu} implies an average contract duration of 6 quarters. While this duration is higher than the 4 quarters advocated in Taylor's (1998) survey, it is approximately consistent with recent estimates of nominal wage stickiness based on more detailed, quarterly data by Barattieri et al. (2010). The inflation indexation factor ω for non-reoptimized wages of 0.5 roughly equals the average proportion of workers receiving cost-of-living adjustments (COLA) in the sample under consideration (see e.g. Hofmann et al., 2010). The elasticities μ^u and μ^{nu} translate into a steady-state markup of 48% and 20%, respectively, for union and non-union workers, implying an average union wage premium of 23.5% in line with the above numbers from Hirsch and Macpherson (2010).

[39] The different calculations ignore 1981 and 1982 as there is no union data for 1982 and the CPS May data for 1981 is not representative of the economy (see Section 3). Since union density estimates only go back to 1964 and were likely to be higher before then, we consider a value of 0.25 for the pre-84 union density as conservatively low.

[40] Specifically, we obtain from the tables in Lemieux et al. (2009a) that performance-pay contracts are about half as likely for union workers than for non-union workers and that the average incidence of performance-pay in the mid 1970s was about 35%. Assuming that the average incidence of performance-pay was 30% for the pre-84 period and 60% for the post-84 period, we can then use average union density rates and the information that performance-pay is half as likely for union workers to compute the values for p^u and p^{nu} reported in Table 6. Our assumption that the proportion of performance-pay contracts approximately doubled is consistent with survey information from Fortune 1000 companies (see Lemieux et al., 2009a for a discussion).

Finally, the elasticity μ is set such that, in combination with the other wage setting parameters, union hours worked are about three times as volatile as non-union hours worked, consistent with the evidence discussed towards the end of Section 3.1. All of the results are robust to reasonable changes in ω, μ^u, μ^{nu} and μ.

For the shock calibration, we let both the technology shock and the preference shock follow an independent AR(1) process

$$\begin{aligned} a_t &= \rho_a a_{t-1} + \varepsilon_{at} \text{ with } \varepsilon_{at} \; iid \; (0, \sigma^2_{\varepsilon_a}) \\ \Delta z_t &= \rho_{\Delta z} \Delta z_{t-1} + \varepsilon_{\Delta zt} \quad \text{with } \varepsilon_{\Delta zt} \; iid \; (0, \sigma^2_{\varepsilon_{\Delta z}}). \end{aligned}$$

The parameters for each process are estimated separately for the subsamples 1953:2-1983:4 and 1984:1-2006:4. For the technology shock process, we use a quarterly measure of total factor productivity constructed by Basu et al. (2006), which controls for variable factor utilization. We convert this measure into logarithms, subtract a linear trend and then estimate ρ_a and σ_{ε_a} by ordinary least squares (OLS).[41] For the preference shock process, we measure Δz_t as the residual from the Euler equation for nominal bonds in (13); i.e. $\Delta z_t = E_t \Delta c_{t+1} - (r^n_t - E_t \pi_{t+1})$.[42] The nominal short-rate in this equation is measured by the 3-month treasury bill rate. Expectations of future consumption growth and inflation are estimated from a bivariate VAR in the two variables, with consumption being measured by real chain-weighted per capita expenditures of non-durables and services and inflation being measured by the growth rate of the GDP deflator.[43] As for total factor productivity, we subtract a linear trend from the obtained series of Δz_t and then estimate $\rho_{\Delta z}$ and $\sigma_{\varepsilon_{\Delta z}}$ by OLS. The point estimates for the pre-1984 and the post-1984 period are given in Table 9.[44] The innovations to both shock processes become less volatile in the post-1984 period. This drop in volatility is, however, much less pronounced for the innovation to the preference shock. Furthermore, the

[41] Substracting a linear trend implies that productivity grows at a deterministic exponential rate, which is consistent with the model.

[42] Alternatively, we could measure Δz_t as the residual from the investment Euler equation in (12). There are two reasons we prefer the bond Euler equation. First, the rental rate of capital in the investment Euler equation has to be inferred from macroeconomic quantities using the firm's capital demand condition in (17). Both the real marginal cost and capital stocks are difficult to measure and thus, we have less confidence in the resulting series for the rental rate of capital than bond prices and inflation, which are directly observable in the data. Second, the investment Euler equation may be affected by investment-specific technology shocks. Primiceri et al. (2006) argue that such investment-specific shocks neutralize a large part of preference shocks, which would lead to a substantially smoother series for Δz_t. These investment-specific shocks do not enter into the bond Euler equation.

[43] Based on Schwarz' Bayesian Information Criterion (BIC), we select a VAR in 5 lags. The different results are robust to alternative lag specifications.

[44] For both sub-periods, the correlation between the innovations is negligible (0.11 and -0.03, respectively). Hence, our assumption that the two shock processes are independent is valid.

preference shock becomes more persistent. As a result, the volatility of the preference shock drops much less and becomes about three times more important relative to the volatility of the technology shock.

4.3 Simulations

The simulation exercise proceeds in four steps. First, we simulate the model with all parameters set to their pre-84 values. Second, we change the shock process calibration to the post-84 estimates and assess to what extent the 'good luck hypothesis' can generate an increase in relative wage volatility. Third, we change $\frac{W^u N^u}{WN}$, p^u and p^{nu} to their post-84 values while keeping the shock processes at their pre-1984 estimates to evaluate the effects of deunionization and higher incidence of performance-pay. Fourth, we simulate the model with both the shock processes and $\frac{W^u N^u}{WN}$, p^u and p^{nu} set to their post-84 values to obtain the joint effect of all changes.

Baseline calibration. The first three columns of Table 10 repeat the U.S. data moments in Table 3. Simulation 1 displays the second moments generated by the model for the baseline pre-84 calibration. As discussed above, the monetary policy parameters are calibrated to match the volatility of output in the data and to get close to the correlation coefficients of labor productivity with output and hours in the data. Despite its relative simplicity, the model does a good job in matching the volatilities of the different labor market variables in the pre-1984 sample. The model also comes reasonably close in generating the high correlation between nominal wages and prices but overpredicts the correlation of wages with output.

Smaller shocks. We now change the calibration of the two shock processes to their post-1984 estimates while keeping all other parameters at their baseline values. As Simulation 2 in Table 10 shows, the smaller volatilities for the two shock processes lead to a substantial fall in output volatility of about 35% as well as a fall in the cyclicality of labor productivity. At the same time, the smaller shock volatilities in the post-1984 period also generate a substantial fall in the volatility of wages with the result that the relative volatility of wages decreases slightly. Hence, while the 'good luck hypothesis' on its own can account for a substantial part of the Great Moderation in our model, it fails to account for the sizable increase in the relative volatility of wages in the data.

To understand these results, it is useful to consider a graphical illustration of the labor market, with the wage setting curve W^S approximating the aggregation of the different optimal wage conditions in (10)-(11) and the curve L^D representing aggregate labor demand in (16).[45] Figure

[45] We provide an explicit description of the loglinearized wage setting conditions for each worker group in the appendix. These conditions can be combined to obtain a linearized expression for the aggregate wage as a function of the aggregate marginal rate of substitution.

3a depicts the response to a positive technology shock. Starting from point A, the technology shock moves labor demand to the right and shifts up the wage setting curve due to the positive income effect on the marginal rate of substitution. The new equilibrium establishes at point B. Smaller technology shocks change the size of these shifts and thus affect the *absolute magnitude* of the reaction in the real wage and labor. However, since the structure behind the two curves remains the same, the *relative magnitude* of adjustments in the real wage and labor remain more or less unchanged.[46] Figure 3b illustrates the effect of a preference shock. The preference shock reduces current consumption, implying a negative income effect that shifts the wage setting curve down. Aside from negligible equilibrium effects on the average markup, the labor demand schedule remains unaffected and thus, the economy adjusts from point A to point B. Similar to the technology shock, smaller preference shocks result in smaller shifts of the wage setting curve. But as long as the slope of this curve remains unchanged, the relative magnitude of adjustments in w and n remains approximately the same. This explains why changes in technology and preference shocks have hardly any effect on the *relative* volatility of wages. By contrast, changes in the relative importance of technology and preference shocks can have important effects on the cyclicality of wages and labor productivity. Technology shocks imply that both wages and labor productivity co-move with hours whereas preference shocks imply exactly the opposite. Hence, when preference shocks become relatively more important, the correlation of wages and labor productivity with hours (and thus output) falls and may even become negative. The graphical illustration suggests that similar conclusions apply for other exogenous shocks that shift either the wage setting curve (e.g. labor supply shocks, government spending shocks) or labor demand (e.g. monetary policy shocks). We confirm this conjecture with robustness exercises in the appendix. By the same arguments, structural changes outside of the labor market (i.e. changes that do not directly affect the nature of wage setting or labor demand) are unlikely to greatly affect relative wage volatility. For example, a change in the responsiveness of monetary policy to inflation (which has been advanced as another possible explanation of the Great Moderation; e.g. Boivin and Giannoni, 2006) would have only a modest effect on the relative volatility of wages since this does not affect the shape of either wage setting curve or labor demand but only by how much they shift in response to shocks (through changes in wage and price markups).

Deunionization and increased incidence of performance-pay. We now reset the calibration of the shock processes to their pre-1984 estimates and instead change the wage share of union workers and the proportion of performance-pay contracts in each sector to their post-84 calibration. As

[46] Our explanation ignores dynamic general equilbrium effects coming through movements in inflation that affect the two curves (see appendix).

Simulation 3 in Table 10 shows, these labor market changes increase the relative volatility of wages by about 55% and reduce the volatility of output by more than 15%. At the same time, the cyclicality of labor productivity and wages increase counterfactually. Robustness exercises in the appendix confirm that similar results obtain for other exogenous shocks.

As before, it is useful to consider a graphical illustration to understand the mechanisms behind these results. Figure 4a depicts the impact of a positive technology shock in a labor market with a relatively steep and a relatively flat wage setting curve. The relatively flat wage setting curve corresponds to a labor market with widespread unionization and little performance-pay where movements in the marginal rate of substitution have little contemporaneous effect on wage setting. A positive technology shock in such a situation leads to a relatively small change in equilibrium wages but a large change in labor and output (point B). As unionization declines and performance-pay becomes more widespread, wage setting increasingly depends on the marginal rate of substitution. The wage setting curve steepens and shifts more with general equilibrium income effects. As a result, the same positive technology shock now implies a much larger equilibrium response of wages relative to the equilibrium response of hours (point C). Furthermore, the correlation of wages with output conditional on technology shocks increases with wage flexibility because the reaction of wages becomes more contemporaneous. Likewise, the conditional correlation of labor productivity with output and hours increases with wage flexibility because productivity shocks affect output proportionally more than hours (due to decreasing returns to scale of hours in production). Figure 4b depicts the impact of a positive preference shock for the same two labor market situations. When unionization is widespread and there is little performance-pay, the income effect of the preference shock is small. Hence, the economy moves to new equilibrium point B, where wages adjust relatively little. Instead, when there is little unionization and performance-pay is widespread, the income effect of the preference shock is larger and the economy ends up at point C where the response of both wages and hours is larger. The larger shifts in the wage setting curve make wages more countercyclical conditional on preference shocks and labor productivity less procyclical (due to decreasing returns to scale of hours in production).

Deunionization, increased incidence of performance-pay and smaller shocks. Finally, we assess the impact of simultaneously changing the union wage share, the incidence of performance-pay and the shock processes to their post-84 calibration values. As Simulation 4 in Table 10 shows, this decreases the volatility of output by a about 45% and increases the relative volatility of wages by over 65%. The combination of 'good luck hypothesis' and greater wage flexibility through deunionization and increased incidence of performance-pay can therefore explain almost the entire drop in output volatility during the Great Moderation and simultaneously generates a substantial increase in relative wage volatility. The model also generates a decrease in the correlations of labor productivity with hours and nominal wages with prices relative to the baseline pre-84 calibration (i.e. Simulation 1). The decrease in correlation of labor productivity with hours is, however, considerably smaller than in the data and the correlation of labor productivity with output barely moves. Moreover, the cyclicality of wages displays a counterfactual increase. Given the small number of shocks in our model, this failure to replicate the different changes in correlations should not come as a big surprise. As discussed above, any additional shock affecting the marginal rate of substitution (e.g. a labor supply shock) that gains in importance relative to the technology shock in the post-84 period would decrease the cyclicality of labor productivity and wages, thus improving the model performance.

In sum, we consider this simulation exercise an instructive and partially successful first step towards a *quantitative* explanation of the great increase in relative volatility of wages. While the increase in relative wage volatility due to deunionization and increased incidence of performance-pay remains below what we see in the data, the exercise highlights how any structural change in the labor market that leads to greater wage flexibility (i.e. wage setting that becomes more sensitive to the marginal rate of substitution) increases the relative volatility of wages and simultaneously reduces business cycle fluctuations.

5 Conclusion

This paper documents that the relative volatility of hourly wages increased by a factor of 2.5 to 3.5 during the Great Moderation. A large part of this increase in relative wage volatility is due to the fact that while output volatility fell by about 60%, the volatility of hourly wages remained roughly constant or increased modestly. CPS microdata reveals that this relative stability in wage volatility applies for many different groups of workers. As a result, the increase in the relative volatility of hourly wages is predominantly due to the increase in relative wage volatility for different groups workers. Compositional changes of the workforce, by contrast, account for no more than 13% of the increase in the relative volatility of hourly wages.

We view these findings as an important challenge for macroeconomic modeling in general and explanations of the Great Moderation in particular. Using a DSGE model, we show that changes in the volatility of exogenous shocks can have a substantial impact on the *absolute* volatility and cyclicality of wages but that these changes on their own have only a small impact on *relative* wage volatility. Hence, the 'good luck hypothesis' that many studies credit as the main driver of the Great Moderation cannot explain the observed large increase in relative wage volatility. Similarly, structural changes outside of the labor market are unlikely to have a large effect on relative wage volatility. This puts the labor market front and center. Motivated by empirical observations, we extend the New Keynesian framework with nominal wage contracts to accommodate unions and performance-pay contracts. We show that deunionization and increased incidence of performance-pay leads to greater wage flexibility in general equilibrium. For reasonable calibrations in line with the U.S. labor market experience, the model generates a substantial increase in relative wage volatility and simultaneously helps account for the Great Moderation.

Our model simulations represent one of the first attempts to provide a *quantitative* assessment – based on calibrations with actual micro-data – of the business cycle effects of structural changes in the U.S. labor market. While deunionization and increased incidence of performance-pay in our model imply a substantial increase in relative wage volatility, much remains to be explained. Given the stylized formalization of unions and performance-pay in our model, this should not come as a surprise. In particular, it is likely that union behavior itself and the nature of performance-pay contracts has changed over the past decades. Building a model that accounts for these changes and can be calibrated from available data remains a challenge for future work.

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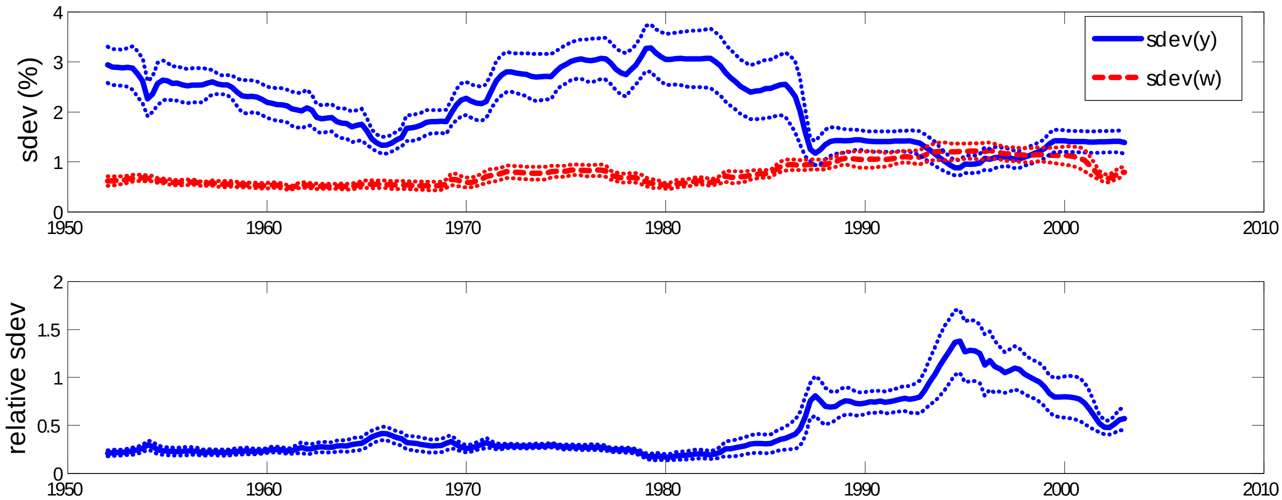

Figure 1: Rolling standard deviations of HP-filtered output and hourly wages (top panel) and rolling relative standard deviation hourly wages relative to rolling (lower panel).

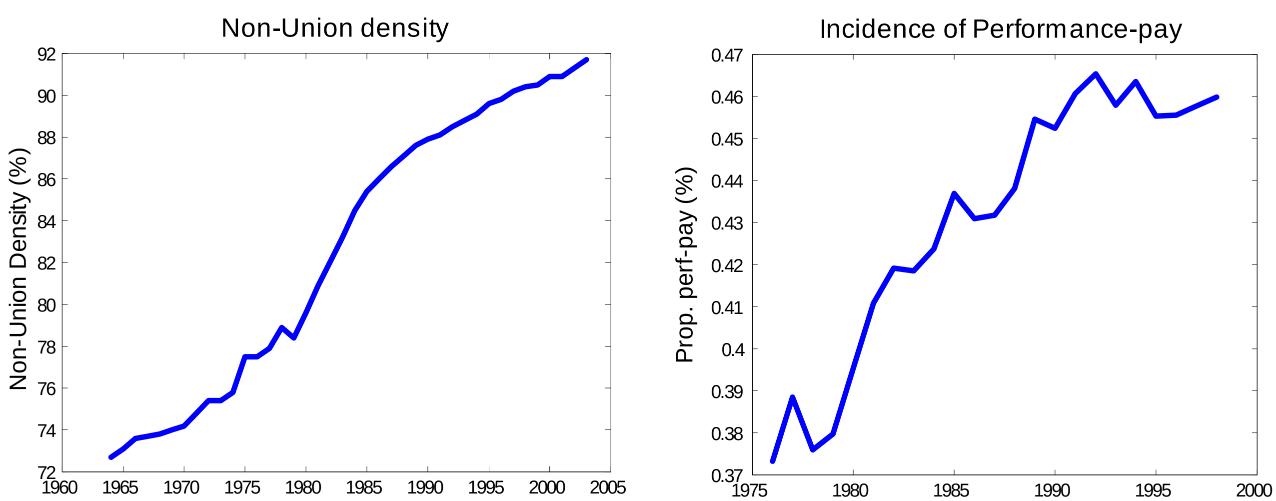

Figure 2: Evolution of non-union density (left panel) from nonfarm business workers and incidence of performance-pay (right panel) in the U.S.

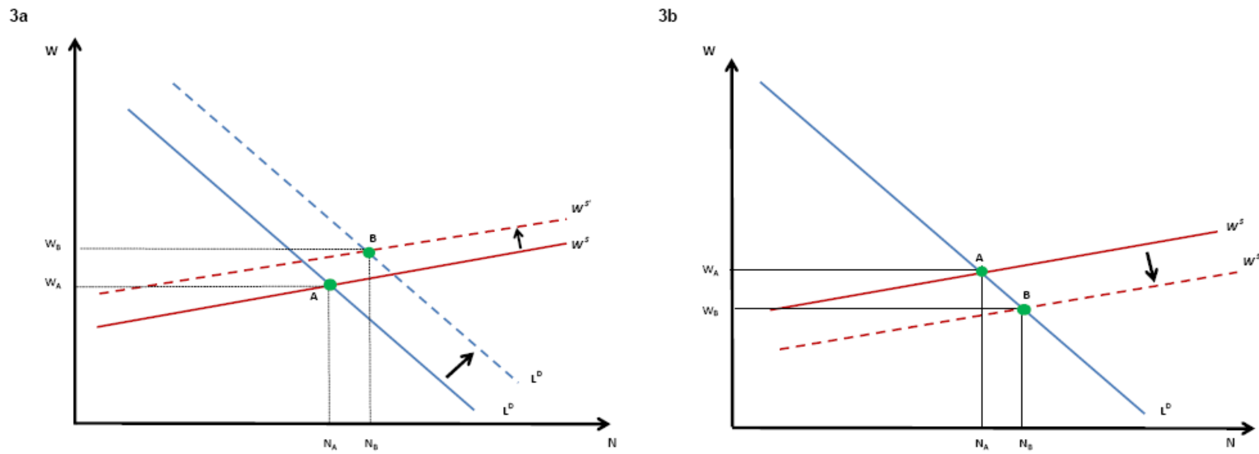

Figure 3: Labor market responses to a positive technology shock (left) and a preference shock (right).

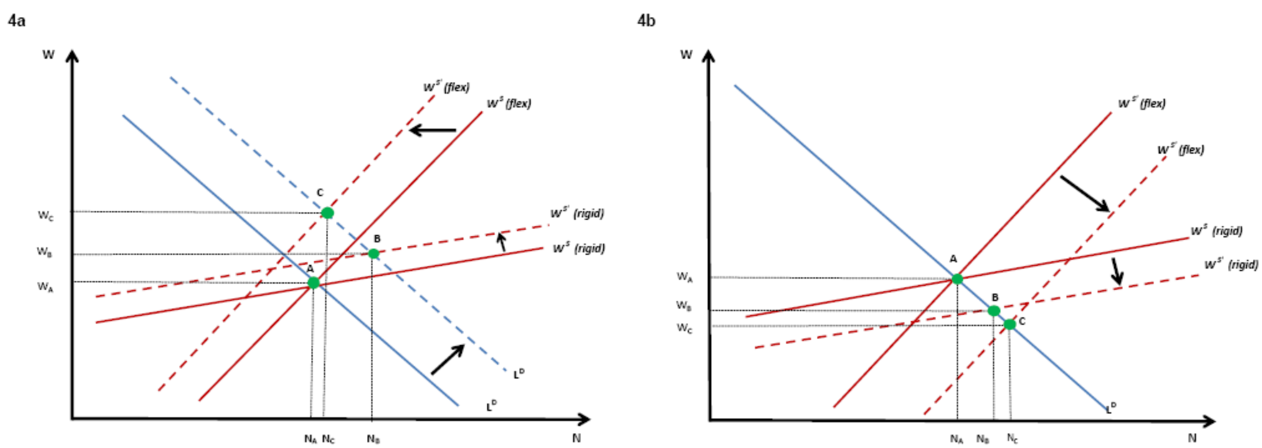

Figure 4: Labor market responses to a technology shock (left) and a preference shock (right) under rigid and flexible wage setting.

TABLE 1
Changes in Volatility

	Standard Deviation				Relative Standard Deviation		
	Pre-84	*Post-84*	*Post/Pre-84*	*p-value*	*Pre-84*	*Post-84*	*Post/Pre-84*
First-Difference							
Output	1.52	0.68	0.45	0.00	1.00	1.00	1.00
	(0.10)	(0.07)					
Wage	0.50	0.68	**1.37**	0.01	0.33	1.00	**3.04**
	(0.03)	(0.07)			(0.02)	(0.12)	
HP-Filter							
Output	2.57	1.28	0.50	0.00	1.00	1.00	1.00
	(0.24)	(0.14)					
Wage	0.63	1.02	**1.62**	0.00	0.24	0.80	**3.33**
	(0.06)	(0.10)			(0.02)	(0.12)	
BP-Filter							
Output	2.50	1.16	0.46	0.00	1.00	1.00	1.00
	(0.26)	(0.11)					
Wage	0.62	0.94	**1.52**	0.00	0.25	0.81	**3.24**
	(0.07)	(0.10)			(0.02)	(0.13)	

Notes : Total sample extends from 1953:2 to 2006:4 with split in 1984:1. Quarterly data. P-values are reported for a test of equality of variances across the two subsamples. Standard errors appear in parentheses below estimates.

TABLE 2
Changes in Volatility

	Standard Deviation				Relative Standard Deviation		
	Pre-84	*Post-84*	*Post/Pre-84*	*p-value*	*Pre-84*	*Post-84*	*Post/Pre-84*
Annual							
Ouput	2.90	1.15	0.40	0.00	1.00	1.00	1.00
	(0.19)	(0.13)					
Aggr. Wage (LPC)	0.60	0.93	**1.55**	0.14	0.21	0.80	**3.89**
	(0.08)	(0.09)			(0.04)	(0.13)	
Aggr. Wage (CPS)	0.63	0.72	**1.14**	0.57	0.22	0.62	**2.86**
	(0.06)	(0.12)			(0.03)	(0.15)	
Quarterly							
Output	2.73	1.28	0.47	0.00	1.00	1.00	1.00
	(0.31)	(0.14)					
Aggr. Wage (LPC)	0.65	1.02	**1.58**	0.00	0.24	0.80	**3.38**
	(0.08)	(0.10)			(0.03)	(0.12)	
Aggr. Wage (CES)	1.11	0.45	**0.41**	0.00	0.41	0.36	**0.87**
	(0.19)	(0.05)			(0.07)	(0.07)	

Notes : Total sample extends from 1964 to 2006 for quarterly data; 1973 to 2006 for annual data; Nonfarm business sector. HP-filtered data. PCE-deflated wages. P-values are reported for a test of equality of variances across the two subsamples. Standard errors computed using GMM and the Delta method appear in parentheses below estimates.

TABLE 3
Changes in Labor Market Dynamics

Moments	*Pre-84*	*Post-84*	*Relative*
$\sigma(n)/\sigma(y)$	0.78	1.15	1.48
	(0.04)	(0.09)	
$\sigma(w)/\sigma(y)$	0.24	0.80	3.37
	(0.02)	(0.12)	
$\sigma(y/n)/\sigma(y)$	0.49	0.59	1.21
	(0.04)	(0.08)	
$\sigma(nomW)/\sigma(y)$	0.37	0.80	2.16
	(0.04)	(0.12)	
$\rho(y,w)$	0.37	-0.14	-0.50
	(0.14)	(0.15)	
$\rho(y,y/n)$	0.65	0.01	-0.64
	(0.07)	(0.14)	
$\rho(n,y/n)$	0.21	-0.50	-0.71
	(0.11)	(0.11)	
ρ(nomW,P)	0.82	0.26	-0.57
	(0.04)	(0.15)	

Notes : All moments are H-P filtered. US data spans from 1953:2 to 2006:4. For correlations, 'Relative' is the difference betw een post- and pre-84, instead of the ratio of post- to pre-84 for standard deviations.

TABLE 4
Changes in Wage Volatility

	Standard Deviation				Relative Standard Deviation		
	Pre-84	*Post-84*	*Post/Pre-84*	*p-value*	*Pre-84*	*Post-84*	*Post/Pre-84*
SKILL / GENDER							
Male unskilled	0.71	0.83	1.16	0.55	0.25	0.72	2.92
	(0.08)	(0.16)			(0.03)	(0.17)	
Male skilled	0.41	1.11	2.71	0.10	0.14	0.96	6.80
	(0.04)	(0.23)			(0.01)	(0.26)	
Female unskilled	0.78	0.73	0.94	0.90	0.27	0.63	2.35
	(0.12)	(0.13)			(0.05)	(0.15)	
Female skilled	1.47	0.84	0.57	0.11	0.51	0.73	1.43
	(0.31)	(0.10)			(0.13)	(0.14)	
SKILL / AGE							
16-29 Unskilled	0.98	1.00	1.02	0.85	0.34	0.87	2.56
	(0.14)	(0.16)			(0.05)	(0.18)	
16-29 Skilled	1.13	1.45	1.28	0.32	0.39	1.26	3.23
	(0.17)	(0.22)			(0.07)	(0.24)	
30-59 Unskilled	0.80	0.76	0.95	0.93	0.28	0.66	2.37
	(0.09)	(0.16)			(0.04)	(0.17)	
30-59 Skilled	0.75	0.94	1.25	0.46	0.26	0.82	3.15
	(0.17)	(0.17)			(0.07)	(0.21)	
60-70 Unskilled	1.35	0.97	0.72	0.15	0.47	0.85	1.81
	(0.20)	(0.08)			(0.08)	(0.12)	
60-70 Skilled	2.65	1.63	0.62	0.03	0.92	1.42	1.55
	(0.48)	(0.19)			(0.20)	(0.26)	
SKILL / EMPL. STATUS							
Hourly, unskilled	0.96	0.89	0.92	0.60	0.33	0.77	2.32
	(0.17)	(0.16)			(0.07)	(0.18)	
Hourly, skilled	1.48	1.48	1.00	0.61	0.51	1.28	2.51
	(0.22)	(0.29)			(0.10)	(0.34)	
Salaried, unskilled	1.21	0.85	0.70	0.24	0.42	0.74	1.76
	(0.21)	(0.08)			(0.07)	(0.12)	
Salaried, skilled	0.44	0.91	2.09	0.14	0.15	0.79	5.24
	(0.04)	(0.19)			(0.01)	(0.22)	

Notes : Total sample extends from 1973 to 2006 with split in 1984; HP-filtered, annual data. Nonfarm business sector.

TABLE 5
Changes in Wage Volatility

	Standard Deviation				Relative Standard Deviation		
	Pre-84	*Post-84*	*Pre/Post-84*	*p-value*	*Pre-84*	*Post-84*	*Pre/Post-84*
SKILL / INDUSTRY							
MinOilGas unskilled	2.40	1.71	0.71	0.16	0.83	1.53	1.85
	(0.32)	(0.30)			(0.15)	(0.34)	
Construct unskilled	1.33	0.96	0.72	0.17	0.46	0.86	1.87
	(0.15)	(0.14)			(0.07)	(0.11)	
Manuf-D unskilled	0.81	1.04	1.27	0.49	0.28	0.93	3.30
	(0.10)	(0.25)			(0.04)	(0.24)	
Manuf-ND unskilled	0.78	1.23	1.57	0.32	0.27	1.10	4.07
	(0.08)	(0.30)			(0.03)	(0.31)	
T&U unskilled	1.13	0.87	0.77	0.39	0.39	0.78	2.00
	(0.16)	(0.13)			(0.07)	(0.19)	
Comm unskilled	2.22	1.29	0.58	0.02	0.77	1.15	1.51
	(0.19)	(0.13)			(0.11)	(0.23)	
Whole T unskilled	1.18	0.89	0.75	0.11	0.41	0.79	1.95
	(0.10)	(0.09)			(0.05)	(0.12)	
Retail T unskilled	1.11	1.01	0.91	0.83	0.38	0.90	2.35
	(0.18)	(0.26)			(0.08)	(0.27)	
FIRE unskilled	1.26	1.01	0.80	0.49	0.43	0.90	2.08
	(0.25)	(0.11)			(0.08)	(0.18)	
Other services unskilled	0.56	0.68	1.21	0.51	0.19	0.61	3.15
	(0.05)	(0.15)			(0.02)	(0.18)	
MinOilGas skilled	5.31	3.86	0.73	0.38	1.83	3.45	1.88
	(0.97)	(0.89)			(0.33)	(1.07)	
Construct skilled	2.38	1.88	0.79	0.35	0.82	1.68	2.05
	(0.27)	(0.37)			(0.12)	(0.46)	
Manuf-D skilled	1.26	1.24	0.98	0.99	0.44	1.11	2.53
	(0.11)	(0.24)			(0.05)	(0.31)	
Manuf-ND skilled	1.53	1.18	0.77	0.21	0.53	1.05	1.99
	(0.15)	(0.07)			(0.07)	(0.15)	
T&U skilled	2.74	2.26	0.83	0.46	0.94	2.02	2.14
	(0.44)	(0.19)			(0.17)	(0.37)	
Comm skilled	4.76	2.17	0.46	0.16	1.64	1.94	1.18
	(1.37)	(0.24)			(0.53)	(0.42)	
Whole T skilled	1.02	1.47	1.44	0.26	0.35	1.31	3.73
	(0.15)	(0.25)			(0.07)	(0.37)	
Retail T skilled	3.04	1.90	0.63	0.04	1.05	1.70	1.62
	(0.29)	(0.22)			(0.10)	(0.26)	
FIRE skilled	0.89	1.25	1.40	0.29	0.31	1.11	3.63
	(0.23)	(0.20)			(0.09)	(0.27)	
Other services skilled	1.52	1.13	0.74	0.19	0.52	1.01	1.92
	(0.15)	(0.22)			(0.07)	(0.29)	

Notes : Total sample extends from 1973 to 2002 w ith split in 1984; annual, HP-filtered data. Nonfarm business sector.

TABLE 6
Relative Volatility Accounting Across Different Decompositions

Decomposition	**Gender/ Skill**	**Age/ Skill**	**Emp Status/ Skill**	**Industry(22)/ Skill**
CPS wage	100.00%	100.00%	100.00%	100.00%
Changing s_i	6.08%	6.49%	12.28%	6.73%
Changing σ*(hourly wages)*^{2}	77.94%	71.05%	70.28%	69.06%
Changing σ*(hours shares)*^{2}	-6.30%	-6.40%	-2.64%	-4.88%
Changing correlations	22.28%	28.86%	20.09%	29.09%

Notes : Total sample extends from 1973 to 2006 w ith split in 1984 (Except for Industry(22)/Education, w hich stops in 2002). HP-filtered data. Nonfarm business sector. Employment status stands for hourly paid or salaried w orkers. Hourly paid w orkers' w ages have been adjusted for the 1994 CPS redesign (see appendix for details).

TABLE 7
Calibration of standard parameters

α	β	γ	δ	$1/\phi$	T/Y	κ	ρ	θ_π	θ_y
0.33	0.99	0.005	0.025	1	0.15	0.05	0.8	2.0	0.3

TABLE 8
Calibration of wage setting parameters

	$\frac{W^u N^u}{WN}$	p^u	p^{nu}	$\frac{1}{1-\xi^u}$	$\frac{1}{1-\xi^{nu}}$	ω	μ^u	μ^{nu}	μ
pre-1984	0.30	0.17	0.32	12	6	0.5	3.1	6	10
post-1984	0.13	0.34	0.64	12	6	0.5	3.1	6	10

TABLE 9
Estimated shock processes

	ρ_a	σ_{ε_a}	$\rho_{\Delta z}$	$\sigma_{\varepsilon_{\Delta z}}$	σ_a	$\sigma_{\Delta z}$
pre-1984	0.9788	0.0094	0.7956	0.0033	0.0549	0.0054
post-1984	0.9738	0.0057	0.8951	0.0020	0.0172	0.0046

TABLE 10
Model Simulations

	US Data			Simulation 1	Simulation 2		Simulation 3		Simulation 4	
				Pre-84 calibration,	*Pre-84 calibration,*		*Post-84 calibration,*		*Post-84 calibration,*	
	Pre-84	*Post-84*	*Relative*	*Pre-84 shock*	*Post-84 shock*	*Relative*	*Pre-84 shock*	*Relative*	*Post-84 shock*	*Relative*
σ(*y*)	2.56	1.28	0.50	2.55	1.65	0.65	2.12	0.83	1.39	0.55
σ(*n*)/ σ(*y*)	0.78	1.15	1.47	0.86	0.93	1.08	0.73	0.84	0.83	0.96
σ(*w*)/ σ(*y*)	0.24	0.80	3.33	0.26	0.25	0.97	0.40	1.56	0.43	1.67
σ(*y/n*)/ σ(*y*)	0.49	0.59	1.20	0.32	0.33	1.02	0.44	1.36	0.43	1.33
σ(*nomW*)/ σ(*y*)	0.37	0.82	2.22	0.29	0.28	0.97	0.42	1.45	0.45	1.53
ρ(*y,w*)	0.36	-0.14	-0.50	0.64	0.65	0.02	0.78	0.14	0.74	0.11
ρ(*y,y/n*)	0.65	0.01	-0.64	0.55	0.36	-0.19	0.76	0.20	0.57	0.02
ρ(*n,y/n*)	0.21	-0.50	-0.71	0.27	0.03	-0.23	0.44	0.17	0.17	-0.09
ρ(*nomW,P*)	0.81	0.28	-0.53	0.63	0.50	-0.13	0.41	-0.22	0.28	-0.35

Notes : All moments are H-P filtered. US data spans from 1953:2 to 2006:4. The 'Relative' column denotes the Post/Pre-84 ratios for standard deviations and the Post-Pre-84 differences for correlations.