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BARGAINING POWER, STRUCTURAL CHANGE, AND THE FALLING U.S. LABOR SHARE

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ABSTRACT

One of the most significant stylized facts in the U.S. economy since the 1970s has been the decline in the share of national income accruing to labor. Many recent studies have sought to explain this trend, with most explanations focusing on structural changes such as deindustrialization, globalization, financialization, rising market concentration, and technological change. We argue that all of these forces primarily operate through a bargaining power channel measured by the cost of job loss, and that the reduction in labor's share of income has been driven by lower bargaining power for workers. Moreover, we contend that business cycle fluctuations in the cost of job loss can help to explain the short-run behavior of the labor share as well. We examine these hypotheses for the United States from 1960-2016. We first estimate the relationship between the cost of job loss and labor's share of income using a bounds-testing approach and find significant negative relationships for both the short and long run. However, the short-run effects are sensitive to the inclusion of policy-related control variables. We then create an index of structural change and estimate regressions of the cost of job loss, finding that increases in this index have both increased the cost of job loss and amplified its volatility over the course of the business cycle. Our empirical analysis therefore supports the hypothesis that the decline in the labor share is driven by decreased labor bargaining power and suggests that structural economic changes and weak economic performance in the U.S. have increased inequality.

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Bargaining Power, Structural Change, and the Falling U.S. Labor Share

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Keywords: labor share, bargaining power, cost of job loss, structural change

JEL Codes: E25, E24, E12

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1. Introduction

One of the most significant stylized facts in the U.S. economy since the 1970s has been the decline in the share of national income accruing to labor. In this paper, we explore the causes of the decline in the labor share in the United States. Building on a theme that consistently appears in the literature, we argue that the reduction in workers' bargaining power has caused the labor share to fall. We empirically measure and analyze the effect of the reduction in workers' bargaining power using multiple measures of the cost of job loss—the one-year income loss due to unemployment—for the 1960-2016 sample, which is determined data availability. Our baseline cost of job loss measure can be seen in Figure 1, along with an index of the labor share.

Figure 1: Wage Share and Real Weekly Cost of Job Loss, 1960-2016

Changes in the cost of job loss over time can be linked to other hypotheses advanced in the literature: structural and institutional changes in the U.S. economy, such as deunionization and deindustrialization; globalization; financialization; increased corporate market power; and technological change and automation. The importance of bargaining power as a channel through which structural changes in economic and institutional relationships can impact the labor share has previously been recognized by Cárdenas and Fernández (2020) and Stansbury and Summers (2020). Most other studies in this literature, discussed in Section 2, show how individual structural changes have directly or indirectly negatively affected the labor share.

The novelty of our approach is twofold. First, we treat structural change as a single process, as opposed to focusing on individual components of this process mentioned above. We view this as a more appropriate approach, both theoretically and empirically, because many of these individual components are closely related in terms of causality and timing, and therefore difficult to differentiate. Second, we use the cost of job loss as proxy for bargaining power. We view this as a more comprehensive measure than alternative variables. To our knowledge, it has not previously been used in analyses of labor share dynamics.

Although the importance of labor bargaining power is intuitive and even explicit in many previous studies of the labor share, we view this as a valuable contribution to the literature. In addition to highlighting the commonality in this expansive literature, focusing on the labor bargaining power channel through which many other factors affect the labor share provides a straightforward theoretical explanation that can likely apply to many different contexts beyond the experience of the U.S. over the past few decades. Admittedly, emphasizing one empirical determinant of the labor share leaves out some of the nuance in the broader literature, wherein various factors that affect the labor share through labor bargaining power are explored in more depth. However, a more parsimonious empirical explanation can be valuable in some contexts, such as multi-equation models examining several interrelationships, wherein degrees of freedom can be limited and including many different explanatory factors for the labor share may not be possible.

Our empirical results suggest that structural changes in the U.S. economy have contributed to a falling labor share and rising functional inequality through a bargaining power channel. We first examine the relationship between the labor share and the cost of job loss. Using a bounds-testing approach, we find a significant negative relationship between the cost of job loss and the labor share in both the short run and long run. Although the significance of the short-run effects is sensitive to the inclusion of control variables representing the aggressiveness of fiscal and monetary policy, the long-run relationship is robust to a number of different specifications.

Given that we find the cost of job loss to be a strong determinant of the labor share, we next examine what drives the behavior of the cost of job loss. In particular, we examine how it is impacted by both the business cycle and long-term structural changes in the economy. To explore the latter, we create a structural index combining proxies for deindustrialization, globalization, financialization, increased corporate power, and technological change. We find that the cost of job loss varies countercyclically and that the business cycle fluctuations in bargaining power are greater in magnitude as the result of these structural and institutional changes. Moreover, we find that higher levels of structural change have increased the cost of job loss overall. Together, these results suggest that structural changes in the U.S. economy and weaker economic performance have reduced the labor share by lowering worker's bargaining power.

2. Literature Review

The literature examining the behavior of the labor share can be divided into two groups: studies identifying factors contributing to the long-term decline in labor's share of national income and studies estimating the relationship between the labor share and economic activity, usually over the course of the business cycle. The former typically highlights explanatory factors that fall into five (non-mutually exclusive) broad sub-categories: structural and institutional changes in labor markets, such as deunionization, deindustrialization, and wage-suppressing policy; globalization; financialization; increased corporate power with monopoly and monopsony market structures; and technological change and automation. Studies in the latter group generally show that the labor share increases with economic activity.

We attempt to synthesize the many wide-ranging elements of this discussion by highlighting the central commonality in all of the determinants of the labor share that have been identified in the literature: they all operate through a channel related to labor bargaining power. The business cycle affects the labor share because labor bargaining power is procyclical, and the factors identified as causes of the long-term decline in the labor share have all, in various ways, lowered labor's bargaining power. Our hypothesis and results are novel because we show that these varied explanatory factors can be comprehensively and parsimoniously captured by movements in the cost of job loss and use this finding to argue that it has caused the decline in the labor share.

Structural and institutional changes in the U.S. economy, particularly in labor markets, have worked to decrease labor's bargaining power and their share of national income. Stansbury and Summers (2020) find that the decline in unionization rates in the U.S. economy reduced workers bargaining power, explaining the decrease in labor's share of income. Using aggregate-, industry-, individual-, and state-level data, they show how declining worker bargaining power since the 1980s shifted rents from labor to capital, resulting in decreases in the labor share and increases in the profit share, in addition to a lower NARIU, which is caused by the employment-incentivizing effects of lower wages. They state:

> The decline in unionization rates and union bargaining power was driven by a combination of institutional factors, which weakened labor law and its enforcement, and economic factors, which increased the elasticity of demand for labor and so weakened workers' ability to bargain for higher wages. Institutional factors included the breakdown of pattern bargaining in the 1980s, the expansion of the number of right-to-work states, and decreasing political support for and enforcement of labor laws. Economic factors that reduced worker bargaining power included increased import competition for manufactured goods and deregulation of transportation and telecoms, both of which reduced firms' abilities to compete while paying high wages. (p. 10)

Based on their argument, the causes of decreased worker bargaining power stem from three sources: institutional change, specifically declining union membership and strike activity; outsourcing or offshoring firm activities to increase profits; and technological change, automation, and low-wage foreign competition resulting from increased globalization. They argue that institutional changes are the main cause of lower worker bargaining power and thus labor's share of income. "Overall, we conclude that the decline in worker power is one of the most important structural changes to have taken place in the U.S. economy in recent decades. Our emphasis on the decline of worker power is justified both by the strength of the direct evidence, and by its ability to provide a unified explanation for a variety of macroeconomic phenomena," including changes in the labor share (p. 7). Additionally, their results "suggest that a large share of the decline in labor rents was a result of a redistribution of rents from labor to capital, rather than a destruction of rents as a result of increased competition or market pressure" (p. 24).

Institutional changes since the 1970s, such as reduced collective bargaining agreements and fewer labor law protections, which weakened labor's bargaining power, are documented in Levy and Temin (2007) and Bivens, Mishel, and Schmitt (2018). Similarly, Mishel and Bivens (2021) highlight key changes in labor markets in recent decades—actively sought by capital and enabled by policymakers—that have dramatically limited labor bargaining power and suppressed wages. These changes include macroeconomic austerity policies that generate excessive unemployment, corporate-driven globalization policies, judicial decisions undermining collective bargaining, weaker labor standards, new contract provisions, and changing corporate structures; and these changes they find to be more important in affecting the labor share than other forces, such as automation. Ramskogler (2021), Brancaccio, Garbellini, and Giammetti (2018), and Ciminelli et al. (2018) argue that the rise of temporary employment contracts, labor market deregulation, and weakened employment protections more generally decrease workers' bargaining power, lowering wage growth and reducing the labor share. Blanchard and Giavazzi (2003) similarly argue that labor market deregulation decreased real wages, but also equilibrium unemployment. The importance of institutional arrangements like collective bargaining agreements is also emphasized by Cárdenas and Fernández (2020). They argue that changes in factors such as collective bargaining arrangements, market structure, and labor movement mobilization during the Francoist developmentalism period in Spain raised the labor share by increasing labor bargaining power. Although they do not test this empirically, focusing instead on the effects of the distributional change, their argument centers on the same mechanisms that we do. Whereas they contend that structural changes in Spain raised the labor share through a bargaining power channel, we argue that a different set of structural changes has caused the opposite to happen in the U.S. Elsby, Hobijn, and Şahin (2013), however, find that changes in bargaining power, measured by the decline in unionization rates across industries, explains only a small fraction of the decrease in the labor share.

Another important structural change in labor markets has been deindustrialization and increasing service sector employment. Beqiraj, Fanti, and Zamparelli (2019) find that the rise of the service sector can help explain the falling labor share in multiple countries. Using U.S. state-level data, Mendieta-Muñoz et al. (2020b) also find support for the deindustrialization hypothesis as real wages become decoupled from productivity and where decreases in the labor share were first seen in goods-producing states, subsequently spreading to all states.

Increases in international trade activity and globalization have had a negative effect of the labor share. Elsby, Hobijn, and Şahin (2013) show that the "decline of the labor share over the last 25 years is largely driven by U.S. producers facing increased import competition" (p. 32). The "import exposure" (p. 27) of industries, where the labor-intensive parts of the production process are offshored to lower-wage foreign countries, lead to lower wages for U.S. workers, decreasing the labor share. Oyvat, Oztunah, and Elgin (2019) argue that an increased international trade lowers the probability of a country experiencing wage-led growth, leading to a decrease in the labor share. Jayadev (2007) and Furceri and Loungani (2018) examine capital account liberalization using a cross-country data set and find a negative effect of the labor share. The rise of import competition from Chinese firms from 1990-2007 served to reduce wages and manufacturing employment in affected sectors (Autor, Dorn, Hansen 2013).

Financialization has also reduced both labor bargaining power and the labor share. Kohler, Guschanski, and Stockhammer (2019) summarize the related literature and identify four different channels through which financialization can affect the labor share, all of which are related to power relationships among classes. These included greater exit options for firms, due to increased capital mobility and the ability for firms to earn profits without hiring workers, higher markups due to financial payments, increased focus on shareholder value and short-term profitability, and rising household debt. Empirically, they find that financial liberalization and financial payments by non-financial corporations have a negative effect on the wage share. Similarly, Pariboni and Tridico (2019) find that financialization has largely been a "redistributive process" (p. 1081) between capital and labor, based on data for 28 OECD countries since 1973, and Stockhammer (2017) finds that financialization, in addition to globalization and the decline of the welfare state, are significant factors in explaining the decline in labor shares in a panel of both developing and developed economies. Ignacio (2015) argues, for the French economy, that increased non-financial firm reliance on profits from the financial sector has lowered labor's bargaining power, which decreases the labor share. Using industry-level data for the U.S. economy, Wood (2017) finds a similar result, and estimates that financialization can account for approximately half of the decrease in the labor share. Dünhaupt (2016) finds that increased dividend payments, which are a result of financialization, depressed the labor share for a panel 13 countries. Financialization is also likely to reduce bargaining power by accelerating deindustrialization and reducing growth of aggregate demand, as Tori and Onaran (2017) show that financialization has had a negative effect on physical investment.

Beginning in the 1980s, corporate power has increased, giving firms more monopoly power in product markets and monopsony power in factor markets. De Loecker, Eeckhout, and Unger (2020) find increasing average markups for large U.S. firms leading to a decrease in the labor share. Autor et al. (2017, 2020) argue that increasing market concentration and the rise of superstar firms drive downward movement in the labor share. Firms that experienced rising market share also had the largest profit share, leading to a decrease in the labor share. Azar, Marinescu, and Steinbaum (2019) estimate that monopsonistic power in labor markets for firms reduces interquartile wages by an average of 17 percent. Although not directly related to firm power, Elsby, Hobijn, and Şahin

(2013) find that the variation in the labor share across industries is larger than the variation in aggregate measures. Falling labor shares in manufacturing, for example, can be offset by increasing labor shares in finance, technology, and health care. Furthermore, since the labor share includes incomes for both payroll and self-employment, where self-employment income is imputed from payroll income, they find that overestimating the decrease in self-employment income artificially decreased the labor share.

Automation and technological change have also worked to decrease labor's share of income (Ray and Mookherjee 2020). Rising levels of automation decrease both the labor share and labor demand, despite increasing productivity (Acemoglu and Restrepo 2019). Ohanian, Orak, and Shen (2021) find that labor share dynamics can be explained by the substitutability between unskilled labor and capital and the complementarity between skilled labor and capital. In a cross-country analysis, Karabarbounis and Neiman (2014) find that the fall in relative prices of investment capital, which induce substitution of capital for labor, decreases the labor share. But Elsby, Hobijn, and Şahin (2013) find that the decrease in the cost of investment capital has not had a significant effect on the decline in the labor share.

There is thus general consensus in the existing literature on the long-term determinants of the labor share that the primary cause underlying the decline in the labor share is the erosion labor bargaining power, due to a number of different factors, although there is disagreement about which are most important. A related and occasionally overlapping strand of the literature examines the relationship between the labor share and economic activity, typically over the course of the business cycle. Although the effect of the labor share on aggregate demand remains a matter of some debate, results throughout the literature generally suggest the presence of "profit-squeeze" effects, wherein the labor share varies procyclically with aggregate demand. (Blecker 2016; Cauvel 2019; Barrales-Ruiz et al. 2020). These effects are likely attributable to changes in labor bargaining power over the course of the business cycle, through channels such as the reserve army of the unemployed (Barrales-Ruiz et al. 2020). Stirati and Paternesi Meloni (2021) demonstrate the importance of this reserve army effect empirically, as their analysis for a number of different countries shows that labor market slack has a negative effect on the labor share.

Some studies in this strand of the literature examine both the short-run relationship between demand and the labor share, as well as changes in this relationship over the long-run. These studies typically identify structural economic changes related to labor bargaining power, such as those discussed above, as explanations for changes in the relationship over the long run. For example, Rada and Kiefer (2015) argue that public policy related to globalization has created a "race to the bottom," as international competition drives down wages and reduces the equilibrium labor share. This effect is amplified by deunionization, the falling price of investment goods, and contractionary monetary policy. Mendieta-Muñoz et al. (2020a) examine how demand, wages, and productivity shocks affect the evolution of the labor share. From 1948-1984, positive productivity, demand, and wage shocks led to the increase in labor's share of national income. But from 1985-2018, negative wage shocks caused by structural and institutional changes in U.S. labor markets were the main driver of the fall in the labor share.

Despite the considerable breadth of the existing literature, which has examined numerous explanatory factors and relationships in both the short and long run, it seems that the behavior of the labor share can be explained quite succinctly: it is driven by changes in labor bargaining power. Cyclical fluctuations in the labor share are driven by profit-squeeze effects that occur because bargaining power varies with demand for labor over the course of the business cycle, while the long-run decline of the labor share in the U.S. is the result of an erosion of labor bargaining power due to various structural changes in labor markets and the macroeconomy. We argue in the subsequent section that a single measure of labor bargaining power—the cost of job loss, the one-year income loss resulting from unemployment—can broadly capture these macroeconomic effects, providing a simple method for empirically capturing much of the variation in the labor share. The cost of job loss is a more comprehensive proxy for labor's bargaining power than other measures such as union power. To our knowledge, this proxy has not been used in the context of explaining the behavior in the labor share. In a two-step process, we show empirically that the cost of job loss and the labor share are negatively related, and that the cost of job loss captures both cyclical changes in bargaining power and the effects of long-run structural changes occurring in the U.S. economy.

3. Data and Measurement

The cost of job loss is a comprehensive proxy for labor's bargaining power because it includes pre- and post-displacement income, labor market dynamics that include dimensions beyond the unemployment rate, and the generosity of income-support programs. It is defined as the total income loss for the one-year period following job loss, expressed as weekly amount. The cost of job loss can be thought of as the "employment rent," or the value of a job to a worker (Schor and Bowles 1987).

Following job loss, a worker loses their pre-displacement income, part of which is replaced by income-support programs, such and unemployment insurance benefits for the fraction of the year they are unemployed. This is defined as a worker's unemployment income. Once rehired, the worker will earn a percentage of their pre-displacement income for the remaining portion of the year. This is defined as reemployment income.

The cost of job loss, in constant 2018 CPI-U dollars, is estimated by subtracting the sum of a worker's weekly unemployment and reemployment income, adjusted for the expected length of unemployment and reemployment, from a worker's weekly pre-displacement income.[1] With a higher cost of job loss, there is a greater value to the employee of keeping their job, which reduces their bargaining power and ability to demand wage increases because of a lower fallback position.

Mathematically, the cost of job loss is defined as:

$$cjl = w - [(UD)w_u + (1 - UD)w_r], \tag{1}$$

where w is pre-displacement income; UD is the average duration of unemployment in weeks, expressed as a percentage of one year; w_u is the total sum of unemployment income; $(1 - UD)$ is the average duration of reemployment in weeks, expressed as a percentage of one year; and w_r is reemployment income, or the income a worker can expect to receive if rehired by another firm (Bowles 1985).[2]

The cost of job loss can be transformed to generate two additional specifications. The normalized cost of job loss expresses the cost of job loss as a percentage of a worker's pre-displacement income, which controls for movements in the pre-displacement wage. To account for the probability of experiencing unemployment, the expected cost of job loss is estimated by interacting the real weekly cost of job loss and the unemployment rate.[3] As will be shown in the next

[1] More discussion and detail on the cost of job loss can be found in Pacitti (2011, 2015).

[2] Our estimate of the cost of job loss assumes that an unemployment spell is less than or equal to one year and that displacement results in unemployment before a new job is acquired, excluding the possibility of a worker making an employer-to-employer transition.

[3] The layoff rate—layoffs as a share of employment—could be used to estimate an additional expected cost of job loss. We experimented with this specification but opted to exclude it in our analysis. When a recession starts, layoffs rise and employment falls, increasing the layoff rate. But the rise in layoffs (the numerator) is amplified by the fall in employment (the denominator). The opposite happens during an expansion. In both cases, the excess variability introduced by this specification does not accurately measure, and likely overstates, labor market dynamics.

section, the five forces that have acted to reduce bargaining power, and thus the labor share, can be more parsimoniously and accurately captured by the costs of job loss, presented in Figure 2.

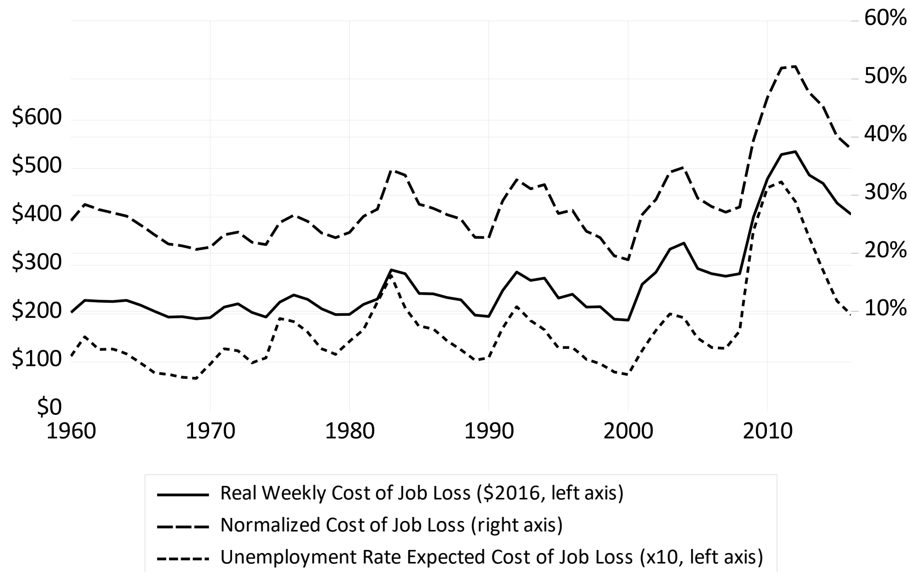

Figure 2: Costs of Job Loss, 1960-2016

To empirically model structural change, we create a weighted index using proxies for the five forces discussed in Section 2—deindustrialization; globalization; financialization; increased corporate power with monopoly and monopsony market structures; and technological change and automation. Deindustrialization is measured inversely by the manufacturing share of total employment, which is calculated as the ratio of manufacturing employment to total private employment. We proxy for globalization using a measure of trade openness for the U.S. economy, defined as the ratio of real imports and exports to GDP. To capture financialization, we calculate the ratio of stock buybacks and dividends to GDP, using the measures for stock buybacks and dividends described in Gruber and Kamin (2017). This measure likely captures two of the channels through which financialization may affect the labor share: increased exit options for firms and increased focus on short-term shareholder value. Markup effects are considered separately, and although we do not directly include a measure of household debt to account for the fourth channel, it is likely to be correlated with this measure of financialization. Our proxy for corporate market power comes from De Loecker, Eeckhout, and Unger (2020), who estimate of the average markup of price over marginal costs for the U.S.[4] Finally, we proxy for technological change using the capital to labor ratio.

Taking the raw values for each of these five variables, we transform them to an index value, ranging between zero (minimum structural change) and one (maximum structural change) (United Nations 2000 and Setterfield and Lovejoy 2006).[5] To generate the weights for each index value, we

[4] Some criticisms of this measure have been raised (Basu 2019; Karabarbounis and Neiman 2018). However, we make use of this proxy because we are unaware of any alternative measures of market power with fewer shortcomings that would cover our entire sample period.

[5] The index value for manufacturing is given by the following formula:

$$\frac{x_{max} - x_t}{x_{max} - x_{min}}$$

For the remaining variables, the index values are given by:

use a principal component analysis, following Bernstein (2014). The weights from this analysis are applied to the index values and we take the average of the five weighted components to calculate our final weighted index. The resulting structural change index is shown in Figure 3.

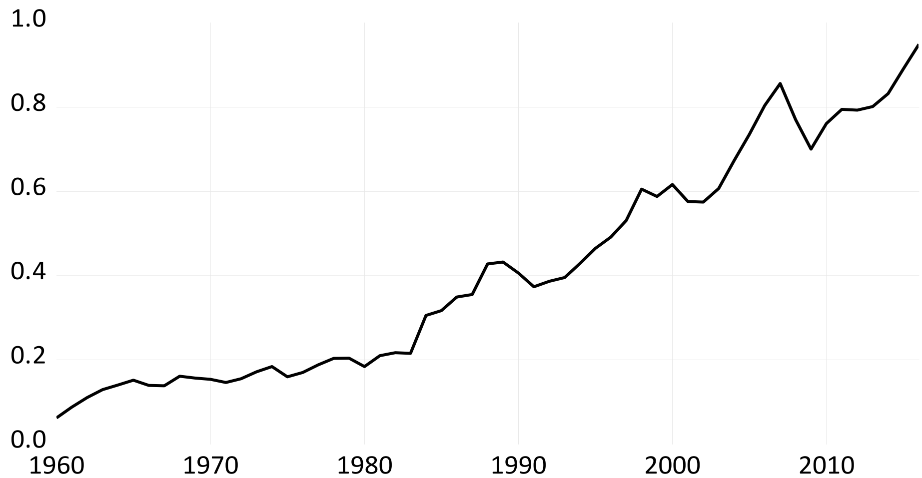

Figure 3: Structural Change Index, 1960-2016

Our analysis also incorporates the labor share, measured using the Bureau of Labor Statistics (BLS) index for the business sector, and the growth rate of real GDP. For comparison to results found using measures of the cost of job loss, we make use of an alternative proxy for labor bargaining power—a measure of union activity showing the number of major work stoppages. Finally, to control for the effects of macroeconomic policy, we construct proxies for the aggressiveness of fiscal and monetary policy. Our proxy for the fiscal policy stance builds upon the methodology of Mahedy and Wilson (2018) for separating the cyclical and structural components of the deficit. They estimate the cyclical deficit as the predicted values of a regression of the primary deficit on the output gap. As we are interested in capturing broader attitudes towards fiscal policy, independent of the business cycle, we use the residuals of this regression—reflecting the structural deficit—as a proxy for the fiscal policy stance. Increases in this series represent greater levels of government spending across the business cycle We employ a similar strategy to proxy for the monetary policy stance, taking the residuals of a regression of the change in the federal funds rate on the output gap. Increases in this variable represent more hawkish monetary policy across the business cycle. However, given that monetary policy is limited by the zero-lower bound, this series should not be interpreted as purely reflecting the attitudes of policymakers, but rather the willingness and ability of policymakers to use expansionary monetary policy.[6, 7]

$$\frac{x_t - x_{min}}{x_{max} - x_{min}}$$

[6] Detailed data definitions, methodology, and sources are included in the appendix. A complete data set is available from the authors upon request, along with results for all unit root tests, diagnostic tests, and unreported specifications discussed in footnotes below.

[7] We log-transformed real GDP as it exhibits exponential growth. We take the log of all other variables as well (with the exception of the fiscal and monetary stance variables used in the bounds testing estimates, which cannot be log-transformed given the presence of negative values) for consistency and because diagnostic tests suggested that preliminary specifications were more econometrically sound when using log-transformed data.

4. Empirical Analysis

Our empirical analysis proceeds in two stages. In the first stage, we estimate the relationship between the cost of job loss and the labor share, first with no control variables and then controlling for the fiscal and monetary policy stance variables.[8] To do so, we use Autoregressive Distributed Lag (ARDL) models and the bounds testing method developed by Pesaran and Shin (1998) and Pesaran, Shin, and Smith (2001), using the helpful guide created by Giles (2015). We selected this approach because we suspected that the cost of job loss and the labor share may be cointegrated, and this method allows us to model the relationship in levels and examine both short-run and long-run dynamics "irrespective of whether the underlying regressors are purely I(0), purely I(1) or mutually cointegrated" (Pesaran, Shin, and Smith 2001). Unlike other error-correction approaches, this method works even if some variables are I(0) and others are I(1). This is particularly useful because unit root tests of most of our variables, including the labor share and cost of job loss series, are not completely conclusive.[9] We therefore use this method to test for cointegration and estimate both short-run and long-run relationships between the labor share and the other variables.

This process involves three steps. First, we estimate an ARDL model in levels, using an algorithm that selects the optimal lag length for each variable by choosing the model that minimizes the Akaike Information Criterion (AIC), subject to the constraint that the lag length cannot exceed four for any variable.[10] We then run various diagnostic tests to ensure a valid econometric specification.[11] Next, we estimate the conditional error correction equation associated with the

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[8] Arguably the government's policy stance also affects the labor share through a bargaining power channel. However, the effects of these control variables may not be fully captured by the cost of job loss.

[9] The augmented Dickey-Fuller (ADF), Phillips-Perron (PP), and Kwiatkowski-Phillips-Schmidt-Shin (KPSS) unit root tests were conducted for all series. The three cost of job loss series all had unit roots according to most of these tests, the two exceptions being the KPSS test for the normalized and excepted cost of job loss. However, these series all appear to have structural breaks around 2007 or 2008 (with 2007 or 2008 as the first year after the break) and modified ADF tests show that the null hypothesis of a unit root can be rejected for all three series if an intercept break (either additive or innovational) is included for either 2007 or 2008. Therefore, we use the log-level of the cost of job loss measures and include an exogenous dummy variable, beginning in 2007 to coincide with the start of the Great Recession, in order to account for the structural break. The same dummy variable is also added to the cost of job loss equations discussed below. Similarly, the standard ADF, PP, and KPSS tests suggest that the labor share has a unit root. However, the null hypothesis of a unit root is rejected in a modified ADF test that allows for a structural break in either 2007 or 2008. The same is true for the fiscal policy stance variable and the union activity series, although in the latter case the PP test does suggest that the series is stationary when no break modeled. The monetary policy stance is the only variable in our labor share estimates that is unambiguously I(0), as all three of the ADF, PP, and KPSS tests suggest that it is does not have a unit root. All three of these tests also show that the first difference of the labor share, fiscal policy stance, union activity, and cost of job loss series are stationary, satisfying the requisite condition for the bounds testing approach that none of the series are I(2).

[10] We use the AIC to determine lag length for these specifications rather than the Schwarz Criterion, which tends to result in the inclusion of fewer lags. Given that there is minimal downside to including more lags in this model, we prefer to use the AIC in order to better ensure that there will be no serial correlation. We deviate from this lag selection process in one case to improve diagnostic test results. In the specification using the baseline cost of job loss measure, the Ramsey RESET test initially suggests misspecification. When adding an additional lag of the dependent variable, we can fail to reject the null hypothesis of no misspecification. Unreported sensitivity tests show that the results are qualitatively robust to using the specification initially suggested by the algorithm instead.

[11] The Jarque-Bera test, Breusch-Pagan-Godfrey test, Breusch-Godfrey Lagrange multiplier test, and Ramsey RESET test were used to test for normality of the residuals, heteroskedasticity, serial correlation, and general misspecification, respectively. No serial correlation up to four lags is found in any specification and the Jarque-Bera test does not indicate any problems with normality of the residuals. The Ramsey RESET test with one fitted term fails to reject the null hypothesis of no specification error for each reported specification. However, the heteroskedasticity tests reject the null hypothesis of homoskedasticity at the 10% level (but not the 5% level) for some specifications (those including the cost of job loss and normalized cost of job loss with the fiscal and monetary policy stance variables, and the specification

ARDL specification.[12] These estimates are used for bounds testing, wherein the null hypothesis is no long-run relationship in levels. If the test statistic exceeds the I(1) critical value the null hypothesis is rejected. If it falls below the I(0) critical value the null hypothesis cannot be rejected, and if it falls somewhere in between the two critical values the results are inconclusive. Coefficients indicating the long-run relationships are calculated based on estimates of the conditional error correction equation.[13] In the final step, a basic error-correction model (ECM) is estimated to examine the short-run dynamics.

We estimate four different specifications, one for each of the three measures of the cost of job loss and one that uses union activity as an alternative measure of labor bargaining power. Each specification includes a constant and a trend. The results are presented in the next section.

In the next stage of our empirical estimation, we examine how both aggregate demand and the five long-term explanatory forces for the declining labor share, as identified in the literature and measured using the structural change index, affect the cost of job loss.

Theory would suggest that structural and institutional changes in the U.S. economy, including deunionization and deindustrialization; globalization; financialization; increased corporate market power; and technological change and automation will lower both pre- and post-displacement wages. Decreased collective bargaining agreements and displaced manufacturing workers who likely find employment in lower-paying service sector jobs; competition with lower-paid foreign and domestic workers; depressed investment in the non-financial economy; increased market power for firms; and a more elastic demand for labor because of automation will, holding constant all other variables unaffected by wages, increase the cost of job loss.

But, as found by Pacitti (2020, 2011), rising costs of job loss have been caused by both secular and cyclical increases in unemployment duration and, by definition, a decrease in reemployment duration, along with lower re-employment earnings, which can persist for decades.[14] The five forces above can be argued to have a similar effect on these variables. Sectoral shifts in economic production, due to deindustrialization, globalization, corporate market power, and automation require the acquisition of costly skills by displaced workers, leading to longer unemployment durations. However, with greater unemployment duration, workers are more willing to accept lower reemployment wages due to the pressures of economic necessity. Together, these forces increase all measures of the cost of job loss, and thus better capture the fall in labor's bargaining power.

It is clear from the data in Figure 2 that the cost of job loss is a cyclical variable, but its relationship to the business cycle appears to have changed over time, as it has risen more during recessions in recent decades. We hypothesize that structural changes in the economy—the five

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including union activity and no control variables other than the structural break dummy). As a result, we use the Newey-West heteroskedasticity and autocorrelation consistent covariance matrix.

[12] As shown in EViews (2020a), a standard representation of an ARDL model with a constant, a_0, and a trend, T, is:

$$y_t = a_0 + a_1 T + \sum_{i=1}^{p} \beta_i y_{t-i} + \sum_{j=1}^{k} \sum_{l_j=0}^{q_j} \gamma_{j,l_j} x_{j,t-1} + \varepsilon_t ,$$

where t indexes time, y is the dependent variable with p lags, the k independent variables, $x_1, \ldots, x_k$, have lags $q_1, \ldots, q_k$, and ε is the error term. The corresponding conditional error-correction representation is:

$$\Delta y_t = a_0 + a_1 T + b_0 y_{t-i} + \sum_{j=1}^{k} b_j x_{j,t-1} + \sum_{i=1}^{p-1} c_{0,i} \Delta y_{t-i} + \sum_{j=1}^{k} \sum_{l_j=0}^{q_j-1} c_{j,l_j} \Delta x_{j,t-1} + \sum_{j=1}^{k} \Delta x_{j,t} + \varepsilon_t.$$

[13] The coefficient for $x_j = \frac{b_j}{b_0}$ (EViews 2020a).

[14] Job-loss induced income losses can be substantial and vary based on the data set and estimation methodology. Couch and Placzek (2010) found earnings losses of 15 percent for six years; Rothstein (2014), Davis and von Wachter (2011), and von Wachter, Song, and Manchester (2009) found losses of approximately 20 percent for as long as 20 years, with losses higher for displacement that occurs during a recession; and Jacobson, Lalaonde, and Sullivan (1993) find losses averaging 25 percent per year.

factors identified in Section 2—have both increased the mean cost of job loss and changed the relationship between bargaining power and the business cycle, leading to larger cyclical fluctuations in the cost of job loss.

We test these hypotheses by estimating ARDL models in which the various measures of the cost of job loss serve as the dependent variable and are regressed on their own lags as well as contemporaneous and lagged values of GDP growth, the structural change index, and the fiscal and monetary policy stance variables.[15, 16] To test whether structural changes have affected the relationship between the cost of job loss and the business cycle, we also interact the structural change index and GDP growth.[17]

This approach is similar to the one we use for the labor share estimates. However, it differs in two ways. First, we do not model these relationships using the bounds testing approach and error correction methods.[18] Second, although we continue to use an algorithm to select the optimal lag length for each variable, subject to the constraint that the lag length cannot exceed four for any variable, here we use the Schwarz Criterion (SC) rather than the AIC, because using the AIC leads to models with large lag lengths that are potentially overfitted.[19] We prefer the more parsimonious specifications suggested by the SC because they enable easier modeling and interpretation of the

[15] We took the first difference of the log GDP series because all three tests of the ADF, PP, and KPSS tests suggested that it had a unit root. All three tests suggested that the structural change index was stationary. These estimates also include the 2007-2016 dummy variable to account for the structural break in the cost of job loss series.

[16] Although we considered various specifications that included the individual components of the structural change index as independent variables in place of the structural change index, we opted against this approach for three reasons. First, as all of the components of the structural change index display similar trends, we suspect that these variables would be highly collinear, making hypothesis testing difficult. Second, using a single variable to capture structural changes enables us to more easily introduce interaction terms, allowing us to test the hypothesis that structural changes have impacted the behavior of the cost of job loss over the course of the business cycle. Finally, our goal is not to test which of the various, likely interrelated, structural changes identified by the literature have been most important. Rather, we aim to explore how structural change, broadly considered, has impacted the cost of job loss.

[17] It is also possible that the structural changes captured by this index have affected GDP growth directly, but testing this is beyond the scope of this paper.

[18] There are two reasons for this. First, preliminary bounds test estimates of the relationship between the structural change index and the baseline cost of job loss (including a constant, a trend, the 2001 and 2007-2016 dummy variables, and no other variables) suggested only weak evidence of cointegration. Although the bounds test F-statistic was above the I(1) critical value associated with the 10% significance level, using the 5% significance level as a threshold, the test statistic was between the I(0) and I(1) critical values, making the results of the bounds test inconclusive. Second, one of the goals of our analysis is to examine how structural changes have affected the relationship between the cost of job loss and the business cycle. Although including interaction terms in an error correction model is possible, this approach would be less straightforward than the one we adopt, and the interpretation of the results would not provide an economically intuitive explanation.

[19] We deviate from this model selection in procedure in one case to improve diagnostic test results. The initial specification suggested by the algorithm appears to have serially correlated errors. We therefore add an additional lag of the dependent variable, because the diagnostic test for this specification no longer rejects the null hypothesis of no serial correlation up to four lags. An unreported sensitivity test shows that results are qualitatively robust to using the specification that minimizes the SC.

interaction terms.[20] The same diagnostic tests used for the labor share estimates are used for these specifications as well.[21]

5. Results

Table 1 presents the estimates for the labor share equation, including long-run coefficients, short-run coefficients, and bounds test results. Eight specifications are reported, two apiece for the three measures of the cost of job loss and two for the alternative bargaining power proxy of union activity. We first estimate the relationship between each of these variables and the labor share with a constant, a trend, and no control variables other than the dummy to account for the structural break. This allows us to test for cointegration between these variables and the labor share. As a sensitivity test, we estimate another specification for each bargaining power variable that controls for the fiscal and monetary policy stance variables.

[20] The lag structure of the interaction terms is based on the lags of GDP growth and the structural change index. We first determine the optimal lag structure for these two variables by running the algorithm for a specification that includes the fiscal and monetary control variables, a constant, and a trend, but no interaction terms. We then create the interaction terms based on the lags of these two variables in that specification. Because the algorithm suggests an optimal specification with one lag of GDP growth and no lags of the structural change index, we include two interaction terms—the product of contemporaneous GDP growth and the contemporaneous structural change index and the product of lagged GDP growth and the contemporaneous structural change index. We then rerun the algorithm with these interaction terms and fixed lags of GDP growth and the structural change index to determine the optimal lag structure for the other variables once the interaction terms are included.

[21] The Ramsey RESET test with one fitted term fails to reject the null hypothesis of no misspecification for each reported set of estimates. The Jarque-Bera test initially rejects the null hypothesis of normal residuals for the specifications using the baseline and normalized cost of job loss measures. We add a dummy variable for the year 2001 to both of these specifications to account for a large outlier, which can be explained by a sharp fall in reemployment wages between 2000 and 2001. The U.S. economy generated higher real wage growth in the latter half of the 1990s expansion, leading to high reemployment wages. With the 2001 recession, a large number of workers exhausted unemployment benefits likely forcing them to accept lower-paying jobs out of economic necessity, driving down reemployment wages (Center on Budget and Policy Priorities 2002). We also include additional dummy variables for 1984 and 1991 in the case of the normalized cost of job loss, as there are large residuals in these years as well. Following the addition of these dummy variables, the Jarque-Bera test suggests that the residuals are normal. Diagnostic tests suggest that errors are heteroskedastic in the normalized cost of job loss equation. As a result, we use the Newey-West heteroskedasticity and autocorrelation consistent covariance matrix for all of our estimates.

	(1)	(2)	(3)	(4)	(5)†	(6)	(7)	(8)
Levels Equation (Long Run) - Dependent Variable: ln Labor Share								
ln Cost of Job Loss	-0.108*** (0.028)	-0.073*** (0.013)						
ln Normalized Cost of Job Loss			-0.112*** (0.027)	-0.065*** (0.012)				
ln Expected Cost of Job Loss					-0.038** (0.017)	-0.024*** (0.007)		
ln Union Activity							0.007 (0.014)	0.028** (0.011)
Fiscal Policy Stance		0.804*** (0.143)		0.922*** (0.157)		0.883*** (0.166)		1.760*** (0.329)
Monetary Policy Stance		-0.005** (0.002)		-0.005** (0.002)		-0.004* (0.002)		-0.013*** (0.004)
ECM Regression (Short Run) - Dependent Variable: Δ ln Labor Share								
Constant	2.443*** (0.457)	3.037*** (0.357)	1.643*** (0.277)	2.739*** (0.325)	1.743*** (0.519)	2.973*** (0.369)	1.653*** (0.435)	1.665*** (0.215)
Trend	-0.000* (0.000)	-0.001*** (0.000)	-0.000** (0.000)	-0.001*** (0.000)	-0.000* (0.000)	-0.001*** (0.000)	-0.000 (0.000)	0.000 (0.000)
Speed of Adjustment	-0.460*** (0.086)	-0.590*** (0.069)	-0.358*** (0.060)	-0.586*** (0.070)	-0.360†† (0.107)	-0.616*** (0.076)	-0.351** (0.093)	-0.366*** (0.047)
Δ ln Labor Share	0.082 (0.105)	N/A^	N/A^	N/A^	N/A^	N/A^	0.244* (0.130)	N/A^
Δ ln Cost of Job Loss	-0.035*** (0.012)	-0.013 (0.010)						
Δ ln Normalized Cost of Job Loss			N/A^	-0.009 (0.010)				
Δ ln Expected Cost of Job Loss					0.000 (0.006)	0.007 (0.006)		
Δ ln Expected Cost of Job Loss (-1)					-0.015* (0.008)			
Δ ln Union Activity							N/A^	N/A^
Δ Fiscal Policy Stance		N/A^		N/A^		N/A^		N/A^
Δ Monetary Policy Stance		-0.001 (0.001)		-0.001 (0.001)		-0.000 (0.001)		-0.002*** (0.001)
Δ Monetary Policy Stance (-1)		0.003*** (0.001)		0.003*** (0.001)		0.002*** (0.001)		0.003*** (0.001)
2007 Dummy	-0.006 (0.005)	0.002 (0.004)	-0.007 (0.005)	-0.001 (0.004)	-0.015** (0.006)	-0.008** (0.004)	-0.021*** (0.007)	0.002 (0.004)
ARDL Lag Structure								
ln Labor Share	2 lags	1 lag	1 lag	1 lag	1 lag	1 lag	2 lags	1 lag
Bargaining Power Variable	0-1 lags	0-1 lags	0 lags	0-1 lags	0-2 lags	0-1 lags	0 lags	0 lags
Fiscal Policy Stance		0 lags		0 lags		0 lags		0 lags
Monetary Policy Stance		0-2 lags		0-2 lags		0-2 lags		0-2 lags
Diagnostics								
ARDL Adjusted R^2	0.935	0.959	0.931	0.959	0.934	0.955	0.909	0.949
ECM Adjusted R^2	0.409	0.638	0.370	0.638	0.405	0.603	0.168	0.548
Bounds Test F-statistic	14.032***	17.004***	17.288***	16.682***	5.531	15.283***	7.073*	14.069***

Table 1. Labor Share Estimates, 1960-2016

Significance levels: ***=1%, **=5%, *=10%. Values in parentheses are Newey-West standard errors. In the case of the bounds test F-statistic, the null hypothesis is no levels relationship, based on the I(1) critical value. For all reported specifications, significance levels for this test are identical regardless of whether the asymptotic or finite sample critical values are used. The significance levels for the speed of adjustment parameter are based on the t-bounds test I(1) critical value.

^Coefficients for Δ ln labor share are only calculated if there is more than one lag of the dependent variable. The first lag of the dependent variable enters the error correction model through the error correction term, for which the speed of adjustment parameter is the coefficient. Similarly, short-run coefficients are only calculated for independent variables if they have more than one lag. Variables with no lags are not included in the estimates, but the contemporaneous values are treated as the sum of the lagged variable and the change in the variable so that they can be properly included in the error correction term (EViews 2020b).
† The addition of the second lag of the expected cost of job loss limits the sample size to 1961-2016.
†† This parameter is not significant based on the I(1) critical values but is above the I(0) critical value for the 10% significance level.

The first specification suggests a strong relationship between the baseline cost of job loss measure and the labor share in both the short and long run.[22] The bounds test F-statistic is above the upper critical value, suggesting cointegration and a long-run relationship. Both the short-run and long-run coefficients are significant at the 1% level, suggesting that the cost of job loss can explain both the short-run and long-run behavior of the labor share. The levels equation coefficient indicates that a 1% increase in the cost of job loss would lead to a roughly 0.11% reduction in the labor share in the long run. This relationship has the expected sign, as reductions in labor bargaining power as measured by increases in the cost of job loss are associated with lower levels of the labor share. The speed of adjustment parameter suggests that about 46% of disequilibrium is corrected within a year. Results are similar when using the normalized cost of job loss measure, although the speed of adjustment is slower and the short-run coefficient on the normalized cost of job loss is not estimated because no lags of this variable are included in the underlying ARDL model.

Relationships between the labor share and either the expected cost of job loss or the union activity variable appear to be considerably weaker. The bounds test F-statistic in the expected cost of job loss specification with no control variables is below both the I(1) and I(0) critical values, suggesting no long-run cointegrating relationship between this measure and the labor share. As such, no inference should be drawn from the coefficients relating this measure and the labor share (in either specification), which are reported only for the sake of completeness.

It is possible that the expected cost of job loss is not a strong measure of labor bargaining power since the unemployment rate may overstate the health of the labor market because it does not account for the duration of a jobless spell, income-assistance program generosity, or reemployment income. Moreover, it is possible that the decline in the unemployment rate overstates the health of the labor market following the Great Recession due to declining labor force participation. Therefore, interacting it with the cost of job loss weakens the results, suggesting the cost of job loss is a more robust proxy for bargaining power when it is included in its original form. In the case of the union activity proxy for bargaining power, there is weak evidence of cointegration, as the bounds test F-statistic is above the I(1) critical value for the 10% level. However, the long-run coefficient is insignificant and no short-run coefficient is calculated because no lags of union activity are included in the ARDL model.

In all specifications including the fiscal and monetary policy control variables, the long-run coefficients on these variables have the expected signs and are statistically significant. A more expansionary fiscal policy stance is associated with a higher long-run labor share, while a more hawkish monetary policy stance is associated with a lower long-run labor share. The contemporaneous short-run coefficient on the monetary policy variable is insignificant in most specifications. The lagged short-run coefficient is significant in all specifications in which it is

[22] All results found using the baseline cost of job loss measure are qualitatively robust to an alternative specification in which the intercept break dummy variable begins in 2001 rather than 2007.

included, but it has an anomalous positive sign. Including these variables increases the speed of adjustment parameter in the error correction model, and the short-run coefficient on either the baseline cost of job loss or normalized cost of job loss is insignificant in these specifications.

One interpretation of this result is that the distributional effects of short-term fluctuations in labor bargaining power, driven by changes in aggregate demand, can be attenuated by government policies, although the long-run equilibrium relationship remains strong. However, we can't rule out the possibility that our method of constructing these policy variables does not completely remove business cycle effects, such as the lagged effects of the output gap on the federal funds rate or the primary deficit. If this is the case, these variables may be capturing some of the effects of demand on distribution that operate through the bargaining power channel, making our bargaining power variable insignificant in the short run. A final possibility is that the aggressiveness of government responses to the business cycle affects the cost of job loss, suggesting collinearity between these variables. We include the fiscal and monetary stance variables in the cost of job loss regressions discussed below to test this possibility. However, given that including these variables increases the adjusted-R^2 values, it is likely that the cost of job loss does not fully capture the distributional effects of monetary and fiscal policy responses.

The long-run coefficients on the baseline and normalized cost of job loss measures are also somewhat reduced when including these control variables, but they remain significant at the 1% level. Conversely, the significance of the union activity measure changes when adding these control variables, as it becomes significant at the 5% level. However, it remains fairly weak, as a 1% increase in union activity is associated with only a 0.03% increase in the labor share. Moreover, the adjusted-R^2 values are higher in specifications including the cost of job loss measures than the proxy for union activity. As such, the results suggest that the cost of job loss may be a better measure for empirically explaining the behavior of the labor share than narrower measures like union activity that capture only one aspect of bargaining power.

The results above suggest that the cost of job loss is a strong determinant of the labor share. In the next step of our empirical analysis, we examine the determinants of the cost of job loss in order to explore the underlying factors that influence the labor share through this bargaining power channel. The cost of job loss equation estimates are shown in Table 2. In these estimates, the individual, "short-run" coefficients represent the direct effects of current (in the case of contemporaneous variables) or past (in the case of lagged variables) changes in the independent variable on the dependent variable in time period t, whereas the "long-run" coefficients represent the cumulative effect, including both the direct contemporaneous and lagged effects (if lags are included) of the independent variable, as well as the indirect effects that it has through lags of the dependent variable in future periods (time t+1 through t+4 in the case of the specification with the baseline cost of job loss, because it includes four lags of the dependent variable).[23]

[23] Because no specification includes lags of more than four years, these effects are "long-run" in an econometric sense, but not necessarily in an economic sense.

	ln Cost of Job Loss	ln Normalized Cost of Job Loss	ln Expected Cost of Job Loss†
Individual Coefficients			
ln Cost of Job Loss (-1)	0.766*** (0.108)	0.687*** (0.077)	0.907*** (0.058)
ln Cost of Job Loss (-2)	0.292* (0.153)	0.461*** (0.087)	
ln Cost of Job Loss (-3)	-0.190 (0.160)	-0.333*** (0.075)	
ln Cost of Job Loss (-4)	-0.097 (0.091)		
Δ ln GDP	-3.220* (1.608)	-3.433*** (1.260)	-12.655*** (2.212)
Δ ln GDP (-1)	-6.743*** (1.606)	-7.075*** (1.447)	-8.252*** (1.456)
ln Structural Change Index	0.109** (0.048)	0.095* (0.050)	0.130 (0.119)
Δ ln GDP * ln Structural Change Index	-1.956** (0.891)	-1.734** (0.672)	-3.897*** (1.340)
Δ ln GDP (-1) * ln Structural Change Index	-2.122** (1.009)	-2.278** (0.884)	-1.980** (0.810)
Fiscal Policy Stance	0.665 (0.632)	0.220 (0.493)	-0.221 (1.088)
Monetary Policy Stance	-0.008 (0.005)	-0.004 (0.004)	-0.029*** (0.008)
1984 Dummy		0.165*** (0.019)	
1991 Dummy		0.152*** (0.018)	
2001 Dummy	0.287*** (0.038)	0.334*** (0.022)	
2007-2016 Dummy	0.017 (0.051)	-0.024 (0.039)	-0.193*** (0.063)
Trend	0.001 (0.003)	0.001 (0.003)	0.001 (0.006)
Constant	1.479*** (0.399)	-0.005 (0.252)	0.813* (0.436)
Long-run Coefficients			
Δ ln GDP	-43.447** (17.681)	-57.137*** (20.825)	-223.665 (145.515)
ln Structural Change Index	0.473* (0.265)	0.517 (0.357)	1.396 (1.937)
Δ ln GDP * ln Structural Change Index	-17.785** (8.234)	-21.815** (8.529)	-62.877 (40.093)
Fiscal Policy Stance	2.898 (2.954)	1.195 (2.641)	-2.365 (11.900)
Monetary Policy Stance	-0.033 (0.024)	-0.020 (0.022)	-0.307 (0.202)
1984 Dummy		0.899*** (0.278)	
1991 Dummy		0.827*** (0.206)	
2001 Dummy	1.252*** (0.407)	1.814*** (0.566)	
2007-2016 Dummy	0.074 (0.205)	-0.132 (0.244)	-2.066 (1.540)
Diagnostics			
Adjusted R-Squared	0.966	0.972	0.955
Sample Period	1963-2016	1962-2016	1960-2016

Table 2. Cost of Job Loss Estimates, 1960-2016

Significance levels: ***=1%, **=5%, *=10%. Values in parentheses are Newey-West standard errors. Long-run coefficients represent the sum of contemporaneous and any lagged coefficients for the independent variable, divided by one minus the sum of coefficients on the lags of the dependent variable. Their significance levels are based on Wald tests of the null hypothesis that the long-run coefficient equals zero, and their standard errors are calculated based on the Delta method.

† The sample period for this specification is 1960-2016. The other two specifications are limited to a sample of 1962-2016 due to the addition of three lags of the dependent variable.

Focusing first on the specification using the baseline cost of job loss measure, the results support the hypothesized relationships. The structural change index is found to be positively associated with the cost of job loss, and both the short-run and long-run coefficients are statistically significant. This finding supports the hypothesis that the various macroeconomic changes captured in the index have reduced labor bargaining power. The long-run coefficient indicates that a 1% increase in the structural change index leads to a 0.47% increase in the cost of job loss. As expected, GDP growth is found to have a negative effect on the cost of job loss, reflecting improvements in labor bargaining power as the economy grows and labor markets tighten.

However, coefficients on the interaction term are significant, indicating that the size of these business cycle effects on the cost of job loss are influenced by structural changes in the U.S. economy. The effect of GDP on the cost of job loss is considerably stronger when the structural change index has a higher value. For example, plugging in the average log-structural change index for the 1960s, a one percentage point decrease in the growth rate of GDP would increase the cost of job loss by about 6.2%, but based on the value of the log-structural change index from 2000-2016, the same decrease in the GDP growth rate would increase the cost of job loss by about 38.1%. These results suggest that structural changes in the U.S. economy have both raised the cost of job loss overall and made it more sensitive to the business cycle. As such, the higher average cost of job loss throughout the 2000s can likely be explained by a combination of direct effects of the various structural changes on labor bargaining power, greater sensitivity of labor bargaining power to the business cycle as the result of these structural changes, and slower average GDP growth.

The fiscal and monetary policy stance variables have the expected signs but are insignificant. This result suggests that the cost of job loss is likely not collinear with these variables in the labor share estimates. However, it is still possible that these variables are capturing some business cycle effects.[24]

We find that the cost of job loss is a strong determinant of the labor share and that the responsiveness of the cost of job loss to the business cycle increases with the structural change index, suggesting that the responsiveness of the labor share to the business cycle should increase along with structural changes in the U.S. economy. The variance of the labor share is 5.7 times larger in the second half of our sample period, relative to the first even though business cycle fluctuations were smaller in magnitude (the variance of GDP growth in the second half of the sample period was only about 49% that of the first half). Our results therefore provide some explanation for the increased volatility of the labor share.

Most of the results found using the baseline cost of job loss measure are robust to using the normalized cost of job loss instead. However, there is one key difference—the long-run coefficient on the structural change index is found to be insignificant. Although the individual, short-run coefficient is weakly significant, suggesting that the structural change has some impact on the

[24] We ran a number of unreported sensitivity tests using the baseline cost of job loss measure. The central results regarding the effects of GDP growth, the structural change index, and the interaction of these two variables are qualitatively robust to using the log-difference (rather than log-level) of the cost of job loss or allowing for asymmetric effects during periods of negative or positive GDP growth through the use of various interaction terms including a dummy variable that takes a value of 1 for years in which GDP growth is negative, all of which were found to be insignificant. The most notable difference in these two specifications is that the cumulative effect of the monetary policy variable becomes significant in the latter and both policy variables become significant in the former, with the expected signs in each case. In two other specifications, one in which the insignificant fiscal and monetary policy stance variables are excluded and one in which the intercept shift dummy variable begins in 2001 rather than 2007, most results are qualitatively robust, but there is one key exception. In both cases, the long-run coefficient for the structural change index is found to be insignificant, although the individual short-run coefficients remain significant at the 5% level.

normalized cost of job loss, the cumulative effect becomes insignificant when accounting for effects through the lagged dependent variable. However, the interaction term remains significant, suggesting that structural changes have still increased the responsiveness of the normalized cost of job loss to the business cycle.

Results are considerably weaker when using the expected cost of job loss. In fact, no long-run coefficients are found to be significant. Even the cumulative effects of GDP growth are insignificant, although the short-run coefficients are significant. This further suggests that interacting the cost of job loss with the unemployment rate, which overstates the strength of the labor market, is not a strong proxy for bargaining power. It can also help to explain why no strong relationship was found between this measure and the labor share, as the underlying factors that we suspect impact the labor share through a bargaining power channel are not strongly related to the expected cost of job loss.

Taken as a whole, the results of our analysis support the hypothesis that both the business cycle and structural changes in the U.S. economy affect the labor share through a bargaining power channel. Bargaining power, as measured by the cost of job loss, is found to be a strong determinant of the labor share in both the short run and long run, although we do find that the short-run effects can be overridden by government policy. The cost of job loss itself is found to be impacted by GDP growth, the structural change index capturing a number of key macroeconomic changes, and the interaction of these two variables. As such, our findings suggest that these macroeconomic changes, along with fluctuations in GDP, influence the functional distribution of income through their effect on labor bargaining power. Our results also suggest that the cost of job loss can better explain the behavior of the U.S. labor share than our alternative bargaining power measure, union activity.

6. Conclusion

We find that structural changes in the U.S. economy since 1960—deindustrialization, globalization, financialization, rising market concentration, and technological change—have had a strong effect on workers' bargaining power, as proxied by the cost of job loss. Our results, based on various specifications of an ARDL model, suggest that these factors, which we capture using an index of structural change, have not only increased the cost of job loss directly, but also increased its sensitivity to business cycle fluctuations. As such, we argue that the combination of these direct and indirect effects of structural changes, along with slower average GDP growth, have caused a reduction in labor bargaining power, as measured by a higher average cost of job loss in the 21st century.

This reduction in bargaining power can explain much of the fall in labor's share of national income. Using the bounds testing method, we find that the cost of job loss is a strong determinant of the labor in both the short and long run. However, the short-run effects are sensitive to the inclusion of variables capturing the aggressiveness of fiscal and monetary policy responses to the business cycle, suggesting that government policy may override the distributional effects of changes in bargaining power in the short run. Our finding that the cost of job loss has become more sensitive to the business cycle as a result of various structural changes in the U.S. economy also provides a potential explanation for increasing volatility in the labor share, although a full exploration of this issue is beyond the scope of this paper.

Our findings provide a potential framework for future research. We argue that subsequent analyses should use comprehensive summary measures, as opposed to individual variables, when looking at long-term structural dynamics because many variables reinforce each other. For example, deindustrialization and deunionization are accelerated by globalization, in addition to automation. Financialization and automation support the rise of large firms, increasing market concentration.

These feedback effects suggest a multivariate measure of structural change could provide a more useful empirical approach to analyzing the changes in labor's share of income. A similar approach is also warranted for bargaining power. Although union membership and the unemployment rate are broad measures of workers ability to claim a share of national income, they omit critical labor market dynamics that also affect bargaining power, such as the secular increase in unemployment duration, the generosity of income-support programs, and the potential for lower reemployment wages. Finally, although our paper does not examine the effects of the falling labor share, its insights might be useful for the vast literature that does, as explicitly modeling this bargaining power channel could improve our understanding of the complex interactions between the labor share and macroeconomic outcomes.

These findings also suggest a large role for policy in shaping economic outcomes, both in terms of equity and efficiency. In the short run, strong fiscal and monetary policy responses can reduce the impact of the business cycle on both output and the income distribution. In the long run, policies that have shaped the U.S. experience with the various structural changes discussed in this paper have resulted in both lower bargaining power for labor and higher inequality. Although an examination of the effects of these changes on output is beyond the scope of this paper, the sharp decline in the labor share and sluggish rates of economic growth in recent decades suggest that these policies have primarily benefitted capital at the expense of labor, rather than creating widely shared benefits. Economists and policymakers should more carefully consider these distributional impacts when examining potential policies. We argue that focusing on bargaining power can be a helpful way to frame these issues. Our analysis suggests that policies that increase workers' bargaining power—such as more expansive income-support programs, guaranteed employment, and pro-union legislation—could be implemented to reduce inequality between capital and labor.

Appendix

The appendix contains data definitions and sources, and supplemental estimates used in the empirical analysis.

Variable	Definition	Source
Real GDP Growth Rate	Log-difference of real Gross Domestic Product (chained 2012 dollars)	BEA via FRED and authors' calculations
Labor Share	Business sector labor share index (2012=100)	BLS Labor Productivity and Costs
Cost of Job Loss	Weekly income loss resulting from job loss ($2016 CPI-U)	Authors' calculations from Pacitti (2011)
Normalized Cost of Job Loss	Weekly income loss resulting from job loss as a percentage of pre-displacement income	Authors' calculations from Pacitti (2011)
Expected Cost of Job Loss	Weekly income loss resulting from job loss multiplied by civilian unemployment rate ($2016 CPI-U)	Authors' calculations from Pacitti (2011)
Union Activity	Number of work stoppages idling 1,000 or more workers beginning in period	BLS via FRED
Nominal GDP	Nominal Gross Domestic Product	BEA via FRED
Primary Deficit	Federal government's current receipts - current expenditures + interest payments	BEA NIPA Table 3.2
Primary Deficit-GDP Ratio	Ratio of primary deficit to nominal GDP	Authors' calculations
Real Potential GDP	Estimate of output at full capacity (chained 2012 dollars)	CBO via FRED
Output Gap	100*(real GDP-real potential GDP)/real potential GDP	Authors' calculations
Fiscal Policy Stance	Residuals of regression shown in TableA2	Authors' calculations
Effective Federal Funds Rate	Weighted average of federal funds rate across transactions	Board of Governors of the Federal Reserve System via FRED
Monetary Policy Stance	Residuals of regression shown in Table A3	Authors' calculations
Manufacturing Share of Employment	Ratio of number of employees in manufacturing industry to number of employees in all private industries	BLS via FRED and authors' calculations
Trade Openness	100*(real exports + real imports)/real GDP	BEA via FRED and authors' calculations
Stock Buybacks	-1*net incurrence of liabilities, corporate equities for nonfinancial corporate business	BEA Integrated Macro Accounts Table S.5.a and author's Calculations
Dividends	Dividends paid by nonfinancial corporate businesses	Board of Governors of the Federal Reserve System via FRED
Financialization	(Stock buybacks/nominal GDP) + (dividends/nominal GDP)	Authors' calculations
Markup	Average markup weighted by market share of sales using time-varying output elasticities	de Loecker et al. 2020
Capital Stock	Capital stock measured at current PPPs (2011 dollars)	Penn World Table, International Comparisons of Production, Income and Prices 9.0
Labor Force	Civilian labor force	BLS via FRED
Capital Intensity	Capital stock / labor force	Authors' calculations
Structural Change Index	Weighted average of principal compents of five structural change variables listed in text	Authors' calculations

Table A1. Data Definitions and Sources

	Primary Deficit/GDP
Output Gap	0.005822*** (0.000893)
Constant	0.007376*** (0.002064)
Adjusted R-squared	0.372204
F-Statistic	42.50121***

Dependent variable: Primary deficit as a share of GDP
Significance levels: ***=1%, **=5%, *=10%
Values in parentheses are standard errors

Table A2. Fiscal Policy Stance

	Change in Effective Federal Funds Rate
Output Gap	0.464468*** (0.081052)
Constant	0.374877* (0.187889)
Adjusted R-squared	0.335712
F Statistic	32.83835***

Dependent variable: Change in effective federal funds rate
Significance levels: ***=1%, **=5%, *=10%
Values in parentheses are standard errors

Table A3. Monetary Policy Stance

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