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Tim Kovalenko, Timo Sauerbier

and Benedikt Schröpf

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The fall and rebound of average establishment size in West Germany*

Tim Kovalenko[1] Timo Sauerbier[1] Benedikt Schröpf[1]

[1] Friedrich-Alexander-Universität Erlangen-Nürnberg, Nuremberg, Germany

Abstract: In West Germany, the average size of establishments declined during the 1990s and started to increase again in the late 2000s, while the employer size wage premium followed the opposite trajectory. In this paper, we show that these two developments are interrelated. More precisely, our results suggest that variations in the employer size wage premiums induced establishments to vary their employment level, consistent with monopsony power on the labor market. Moreover, our regional analyses show that average establishment size correlates positively with GDP per capita. We rationalize these findings with a heterogeneous firms model with monopsonistic competition in the labor market, stemming from the household's love-of-variety preferences for employers. Both empirics and theory reveal that higher size wage premiums decrease average establishment size by downsizing incumbent establishments and triggering the entry of small establishments, thus also negatively affecting aggregate productivity.

Zusammenfassung: In Westdeutschland ging die durchschnittliche Betriebsgröße in den 1990er Jahren zurück und begann in den späten 2000er Jahren wieder zu steigen, während die Betriebsgrößenlohnprämie umgekehrt dazu verlief. In diesem Papier zeigen wir, dass diese beiden Entwicklungen miteinander zusammenhängen. Unsere Ergebnisse legen nahe, dass Schwankungen der Betriebsgrößenlohnprämien Betriebe dazu veranlassten, ihr Beschäftigungsniveau zu variieren, was auf Monopsonmacht auf dem Arbeitsmarkt hindeutet. Außerdem zeigen unsere regionalen Analysen, dass die durchschnittliche Betriebsgröße positiv mit dem BIP pro Kopf korreliert. Wir erklären diese Befunde anhand eines theoretischen Modells mit heterogenen Firmen und monopsonistischen Wettbewerb auf dem Arbeitsmarkt, der sich aus love-of-variety-Präferenzen der Haushalte für unterschiedliche Arbeitgeber ergibt. Sowohl Empirie als auch Theorie zeigen, dass höhere Betriebsgrößenlohnprämien zu einer Verringerung der durchschnittlichen Betriebsgrößen führen, indem bestehende Betriebe ihre Beschäftigung reduzieren und der Eintritt kleinerer Betriebe in den Markt gefördert wird, was sich ebenfalls negativ auf die aggregierte Produktivität auswirkt.

Keywords: Establishment Size, Size Wage Premium, Productivity, Labor Market Power, Germany
JEL classifications: E24, J31, J42, L25

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1 Introduction

The average establishment size (in terms of employees) is an important outcome of an economy. This is mainly because the average establishment size of a country correlates positively with productivity (growth) (Pagano & Schivardi, 2003; Bento & Restuccia, 2021). Hence, economies with larger firms tend to be more productive. On the micro-level, this is consistent with the existence of scale effects that provide productivity gains for larger establishments. Despite its relevance, careful assessments of the causes and consequences of variations in the average establishment size are rather scarce in the literature.

In this paper, we document a fall in the average establishment size in the 1990s and early 2000s and a rebound in subsequent years in West Germany and study the causes and consequences of this evolution. Two main components have contributed to the fall and rebound of the average establishment size in West Germany. First, the average size of incumbent establishments declined substantially in the 1990s and started to recover in the mid-2000s. Second, we observe a marked influx of new establishments in the 1990s, which peaked at the turn of the millennium and dramatically declined afterwards. As new entries tend to be smaller than incumbent establishments, they reduce the average establishment size.

To study the macroeconomic consequences of this evolution, we exploit regional variation within West Germany, thereby establishing a marked and statistically significant relation between average establishment size and labor productivity. According to our estimates, the fall in the average establishment size in the 1990s and early 2000s was associated with a decline in GDP per capita by roughly to 2,000 euros. In contrast, the rebound of the average establishment size was associated with an increase in GDP per capita by around 3,000 euros. These sizeable estimates render the investigation of the dynamics in the average establishment size an important topic.

Why did the average establishment size in West Germany vary this much during the last thirty years? In this study, we develop an explanation based on the evolution of the establishment size wage premium, which followed the opposite trajectory to the average establishment size. Hence, this premium substantially increased during the 1990s and early 2000s and declined afterwards (see Colonnelli et al., 2018; Lochner et al., 2020). Following Bachmann, Bayer, et al. (2022), we interpret the size wage premium as an upward-sloping labor supply curve to the firm and, hence, as an indicator of monopsony power in the labor market. Therefore, we argue that an increasing size wage premium induces establishments to reduce their employment levels since it incentivizes establishments to stay or become small to avoid increasing wages and maximize profits (and vice versa if the size wage premium decreases) (see Bachmann, Bayer, et al., 2022). Exploiting regional variation, we find a statistically significant and sizeable negative relation between average establishment size and the size wage premium (i.e., the level of monopsony power). Furthermore, our empirical analyses reveal that regions with higher size wage premiums experience a greater influx of new establishments. These findings are consistent with the identified components of the dynamics in the average establishment size.

To rationalize our findings, we propose a dynamic general equilibrium model with heterogeneous firms and monopsonistic competition in the labor market, which largely builds on Ghironi and Melitz (2005), Jha and Rodriguez-Lopez (2021), Berger et al. (2022) and Colciago and Silvestrini (2022). In this model, employers are imperfect substitutes from the household's perspective, which ties workers to their employers and results in monopsonistic competition in the labor market (reflected by a size wage premium). Consequently, employers are endowed with a certain degree of wage-setting power and use it to maximize profits, thereby reducing their size (and wages). Using this model, we can replicate the empirical finding of a declining average establishment size when monopsony power and, therefore, the size wage premium increases. Moreover, we can pin down two channels contributing to the decline in the average establishment size. A higher size wage premium induces incumbent firms to reduce their employment as this decreases the wage for the marginal worker. Furthermore, a rise in monopsony power increases firm profits and, therefore, more new (or previously inactive) firms enter the market, which reduces the average firm size further. The model also predicts a decrease in aggregate productivity in response to an increase in monopsony power as the weight of productive firms decreases. In sum, these model results are consistent with our empirical findings.

This paper contributes to the literature in three ways. First, we confirm the result of Pagano and Schivardi (2003) and show that the average establishment size is an important factor for the aggregate economy. More precisely, we demonstrate that regional average establishment size is positively associated with regional labor productivity in West Germany. Second, we establish an economically meaningful relation between the average establishment size and the size wage premium. This is in line with the findings of Bachmann, Bayer, et al. (2022), who argue that average establishment size differences between East and West Germany are largely driven by differences in their size wage premiums. However, we are the first to establish this link within West Germany. Third, we present a simple dynamic macro model with monopsonistic competition in the labor market. The model can be seen as a dynamic version of the closed economy model by Jha and Rodriguez-Lopez (2021) with the micro foundation for monopsonistic competition taken from Berger et al. (2022). By simulating the model, we demonstrate the connection between the degree of monopsony power, the average firm size and productivity.

From a policy perspective, our results yield important insights. A key implication of our study is that productive firms stay or become inefficiently small through monopsony power, thereby reducing aggregate productivity. Hence, policymakers should be interested in making the labor market more competitive. While our model does not provide a straightforward suggestion, one possible strategy could be to increase employees' mobility by reducing commuting costs or fostering working-from-home policies. This can weaken employees' ties to their employers and, thus, reduce monopsony power. Moreover, our analysis suggests that the influx of new establishments in the 1990s was partly driven by an increase in monopsony rents. This may have induced new and less productive firms with previously unprofitable business models to enter the market, thereby reducing aggregate productivity. While strong business dynamics have clear economic upsides, such as fostering innovation and competition, this potential drawback should be kept in mind when designing policy measures that aim to facilitate new business formation.

The paper is organized as follows. In section 2, we relate our paper to the existing literature. Section 3 describes the data and data preparations. The evolution of the average establishment size and its components is illustrated in section 4. Section 5 establishes the association between average establishment size and labor productivity, while section 6 links the size wage premium to the average establishment size. In section 7, we present the dynamic model that rationalizes the empirical findings. Section 8 concludes.

2 Related literature

In this paper, we study the economic relevance of variations in the average establishment size of an economy and establish a link between the average establishment size and the size wage premium. Therefore, our paper relates to the literature on determinants and consequences of the average firm size and the literature on employer size wage premiums and monopsony power.

Previous papers have shown that average firm size correlates positively with productivity (growth) (Pagano & Schivardi, 2003; Bartelsman et al., 2013) and GDP per capita (Bento & Restuccia, 2021). Similarly, Guner et al. (2008) and Gourio and Roys (2014) argue that size-dependent policies reducing the average establishment size also reduce output and output per worker, respectively. Moreover, Bento and Restuccia (2021) exploit variation across countries and sectors and find that the average establishment size correlates positively with external financing (which serves as a proxy for financial constraints) and openness to trade and negatively with firing costs. Their preferred explanation is that the average establishment size is determined by the misallocation of factors across heterogeneous establishments. Hence, variation in average establishment sizes across countries stems from variation in the extent of misallocation across countries, a view in line with previous work of Bento and Restuccia (2017), Hopenhayn (2014) and Hsieh and Klenow (2014). In a recent paper, Bachmann, Bayer, et al. (2022) propose an alternative mechanism of misallocation that distorts the average establishment size, namely the downsizing of establishments due to an increasing size wage premium. The rise in this premium gives establishments an incentive to stay small, which dampens aggregate productivity in the economy. Hence, an increasing size wage premium generates a misallocation of labor from productive to less productive establishments. Bachmann, Bayer, et al. (2022) argue that differences in the size wage premiums between East and West Germany largely explain persisting productivity differences between the two regions. In this paper, we follow their line of reasoning. However, we apply it to the fall and rebound of the average establishment size in West Germany.

Both the evolution and meaning of the size wage premium have been studied by a wide range of papers. Bloom et al. (2018), for instance, document a decline in the size wage premium in the US since the 1980s. In contrast, Colonnelli et al. (2018) and Lochner et al. (2020) find an increase in the size wage premium in the 1990s and early 2000s in Germany and a downward trend afterwards. Even though we interpret the size wage premium as an upward-sloping labor supply curve to the firm and, accordingly, as an indicator of monopsony power in the labor market, this is by no means the only proper interpretation. For instance, a positive size wage premium can potentially stem from the selection of skilled workers into large establishments. Further prominent explanations are that a size wage premium reflects compensating differentials if workers dislike working in large firms or efficiency wages (to avoid shirking), as monitoring in larger firms is more costly. In addition, rent sharing in large and productive firms could also be reflected in a positive relationship between size and wages (see Colonnelli et al., 2018; Green et al., 1996).

We interpret the size wage premium as an upward-sloping labor supply curve to the firm for three reasons. First, other measures of monopsony power have evolved similarly over time in Germany (see Bachmann, Demir, & Frings, 2022; Hirsch et al., 2018). Second, our evidence suggests that the size wage premium is negatively associated with the average establishment size, which, in turn, is positively associated with productivity. Given these pieces of evidence, we think that the interpretation as monopsony power is the most credible explanation since efficiency wages and compensating differentials are profit-maximizing wage-setting regimes, where firms would not have to re-optimize and reduce their size. Also, both strands of explanation cannot rationalize the influx of new establishments. Moreover, if the extent of rent sharing increases in size, firms would not be incentivized to reduce their workforce. Third, we present suggestive evidence on the evolution of concerns about job security and quit rates implying that workers' reluctance to change their jobs rose during the 1990s. This is a source of monopsony power since it would tie workers more strongly to their employers, thereby increasing the labor market power of the firm.

The seminal work of Manning (2003) popularized the view that labor markets should be regarded as imperfectly competitive and are characterized by certain degrees of monopsony power. A necessary implication of monopsony power is that workers will accept certain wage markdowns (relative to their marginal productivity) since they cannot perfectly substitute between employers. This translates into labor market power of the firm and is reflected by an upward-sloping labor supply curve. In modern approaches, monopsony power mostly stems from either the existence of search costs (making job changes less attractive) or heterogeneous preferences for particular employers (which workers pay for by waiving a fraction of their marginal productivity) (Manning, 2021). In models where search frictions are the source of monopsony power, it is usually measured as the difference between the wage elasticity of recruitment and the wage elasticity of the separation rate (Manning, 2003; Hirsch et al., 2018, 2022; Bachmann, Demir, & Frings, 2022). However, we follow approaches that measure monopsony power with the size-wage relation of firms, such as Bachmann, Bayer, et al. (2022) or Green et al. (1996).

From the modelling perspective, we borrow mainly from three previous studies. We build a dynamic model with heterogeneous firms along the lines of Ghironi and Melitz (2005), where firm heterogeneity manifests in productivity differences between firms. These differences in productivity imply differences in firm size. To incorporate the observed size wage relation, we model the labor market with monopsonistic competition following Berger et al. (2022). The key idea is that the disutility of work is lower if the household has a large set of potential employers to choose from, implying that employers are no perfect substitutes. Hence, each employer has some labor market power (reflected by a size wage curve). Furthermore, as in Jha and Rodriguez-Lopez (2021), we allow for endogenous firm entry in the model. In this way, we can establish an economically meaningful channel for the observed correlation of the size wage premium and establishment entries.

3 Data

We use two sources of data for the empirical analyses. Most of the analyses are based on the Establishment History Panel (BHP) provided by the Institute for Employment Research (IAB). It is a comprehensive dataset containing a 50% random sample of all establishments in Germany with at least one employee subject to social security contributions. The establishments are observed yearly on June 30th. Thus, it is a panel dataset. The data source for the BHP is the Employee History (BeH), which is based on the notification procedure for health, pension and unemployment insurance. All employers are obliged to notify all employees who are subject to social security contributions, making it a highly reliable dataset. Since civil servants and the self-employed are not subject to social security contributions, they are not recorded in this dataset (see Ganzer et al., 2021).

The BHP provides information on average gross daily wages and the number of full-time and part-time employees in each establishment. However, there is no information on hours worked (per worker). Since we need comparable wage information to estimate the size wage premium, we base our analyses on full-time employees. Hence, we generally exclude all establishments without at least one full-time employee. Moreover, the BHP provides information on several establishment characteristics such as location on the county level, economic sector (according to the German Classification of Economic Activities 1993) and workforce composition (in terms of gender, nationality, skill level, age and occupation).

The IAB provides an extension file to the BHP to identify establishment entries and exits, based on the classification procedure of Hethey-Maier and Schmieder (2013). Merely relying on new or not recurring establishments ID numbers may lead to misleading results since they might reflect ID changes, spin-off establishments (e.g., domestic outsourcing) or reconstructions (alongside true entries or exits). To identify entries and exits, Hethey-Maier and Schmieder (2013) use clustered inflows and outflows of employees. The idea is the following: if a significant fraction of worker inflows to an establishment with a new ID is from another establishment, likely, an ID change (if the parent establishment ID does not exist any more) or a pulled spin-off (if the parent establishment ID continues to exist) is observed. If, on the other hand, the inflows of an establishment with a new ID come from various other establishments, this establishment is classified as a new production entity. In particular, Hethey-Maier and Schmieder (2013) identify the following cases as reasons for the occurrence of a new ID: new establishment, pushed spin-off (predecessor establishment exits), pulled spin-off (predecessor establishment continues to exists) and ID change. Furthermore, they identify the following cases for the disappearance of an ID: exiting establishment, pushed spin-off, take-over/reconstruction (for the exact definitions, see Hethey-Maier & Schmieder, 2013).

In the following analysis we differentiate between incumbent, newly entering and exiting establishments. We define incumbents as establishments that did not enter or exit in a given year. Additionally, ID changes, takeovers/reconstructions and pushed and pulled spin-offs are counted as incumbent establishments. We classify new entries and exits according to the residual categories in the definitions of Hethey-Maier and Schmieder (2013). Hence, both types of spin-offs are defined as incumbents since we consider these types of changes as means of adjusting the establishment size (e.g., establishing a pulled spin-off may be more cost-effective than growing large and paying higher premiums). However, other classifications of incumbents, entries and exits do not change the implications of our analyses (results upon request).

To estimate the size wage premiums, we employ two approaches. First, we follow the recent literature and rely on AKM establishment effects (see Bloom et al., 2018; Lochner et al., 2020).

These are establishment-specific wage premiums, adjusted for observable and unobservable workforce characteristics (Abowd et al., 1999). Bellmann et al. (2020) provide estimates for the German labor market by performing the following estimation:

$$\log(w_{it}) = \alpha_i + \psi_{J(i,t)} + X_{it}'\beta + e_{it}. \quad (1)$$

Where w_{it} is the real daily wage of individual i in year t, α_i the individual fixed effect, $\psi_{J(i,t)}$ individual i's establishment in period t and X contains variables controlling for individual i's observable characteristics (a set of year dummies as well as quadratic and cubic terms in age fully interacted with educational attainment). The establishment fixed effect $\psi_{J(i,t)}$ is interpreted as establishment j's wage premium paid to all employees independent of the employees' observable and unobservable characteristics. Thus, we will use $\psi_{J(i,t)}$ as the establishments' payment measure adjusted for these worker characteristics. Since $\psi_{J(i,t)}$ can only be identified by job changers, Bellmann et al. (2020) estimate equation (1) for the five intervals 1985-1992, 1993-1999, 1998-2004, 2003-2010 and 2010-2017.

In our second approach, we rely on establishments' average wage levels that is provided in the BHP data. We use the average gross daily wage of full-time employees as a proxy for an establishment's wage level. We employ an imputed version of this wage variable, implemented by Card et al. (2015), since initially earnings are right censored as they are only reported up to the upper limit for statutory pension insurance contributions. Consequently, a substantial share of the wage information (roughly 10 %) is censored at the top (Ganzer et al., 2021).

Since both the size structure and the size wage curves of East German establishments substantially differ from those of West German establishments, we include only West German establishments in our analyses.[1] In addition, many labor market studies find different results in West and East Germany, rendering a comparison of these regions difficult (see Schnabel, 2016). Moreover, in the estimations of the size wage premium and in all analyses based on it, we exclude

[1]For more details on these differences, see Bachmann, Bayer, et al. (2022).

establishments with less than ten employees as very small establishments might follow less strategic wage-setting and personnel adjustment regimes than larger establishments. Furthermore, they are exempt from dismissal protection. Also, we exclude non-private establishments and the farming and mining sectors.

To relate average establishment size to, e.g., productivity, we aggregate the establishment size information on the regional level. Hence, we exploit regional variation to establish the economic relations described above. In the main specifications, we analyze data on the county level. However, we also use local labor market regions based on commuting zones as robustness checks. We employ two definitions of local labor markets: the definition according to "joint task of the federal government and the states for the improvement of regional economic structures" (GRW) (see Breidenbach et al., 2018) and the definition of Kosfeld and Werner (2012). Both use a similar approach based on commuting zones. A major difference between these two definitions is the maximum commuting time allowed within a local labor market region. The 324 West German counties correspond to 203 local labor markets according to GRW and 108 local labor markets according to Kosfeld and Werner (2012). Thus, analyzing labor market regions instead of counties reduces the number of observations. However, these labor market regions have the advantage that they are more likely to represent self-contained labor markets and, hence, are more likely to reveal general equilibrium effects.

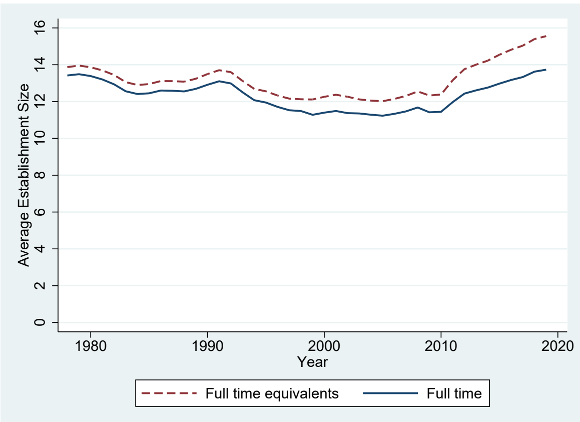
Our second data source is the German national accounts (VGR). We use the national accounts provided by the West German federal states. These data include information on GDP per capita, employees and hours (the latter only from 2000 onwards) on the county level. We deflate these data with the national CPI since no regional price indexes are available (the base year is 2015). The data spans from 1992 to 2019 but since the year 1993 is missing, we effectively use the period between 1994 and 2019. Moreover, due to reforms of county borders, we observe counties in the federal state Saarland from 1996 onwards and counties in Lower Saxony from 2000 onwards.

4 Establishment size distribution over time

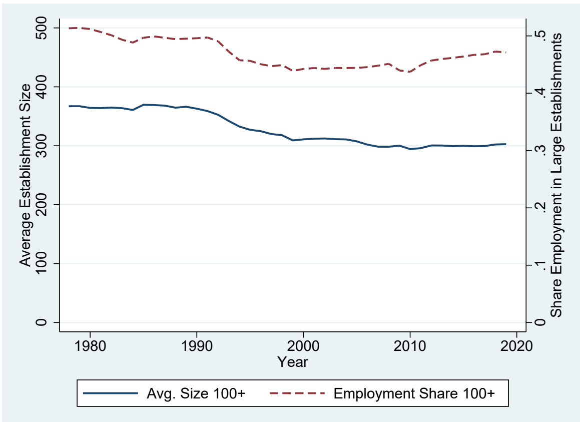
This section describes the evolution of the establishment size distribution over time in West Germany. Starting with the average establishment size displayed in Figure 1 (left panel), we see a fairly stable average size of about 13 full-time employees per establishment in the 1980s. During the 1990s, the average establishment size declined substantially by a total of more than one-ninth. This declining trend lasts until the early 2000s when the size distribution plateaus at over eleven employees per establishment. Starting in the early 2010s, we observe a reversed trend as the average establishment size rises sharply until 2019 (the last data point available to us). Accordingly, the average establishment size rose by roughly one-fifth between 2000 and 2019. Thus, over 30 years, we document a fall and rebound of the average establishment size in West Germany. We can show that this is not due to an increasing share of part-time employment by depicting the average establishment size based on full-time equivalents (where we assign the weight 0.5 to a part-time worker due to the lack of data on hours worked). As can be seen, the evolution follows the same path.

Figure 1: Average establishment size

(a) Avg. est. size considering full-time workers and full-time equivalents



(b) Avg. est. size for establishments with ≥ 100 employees and their employment share



The decrease in size is especially pronounced in large establishments, which we demonstrate in the right panel of Figure 1. The average size of establishments with at least 100 full-time employees fell from about 370 to 300 during the 1990s (a decrease by about 20%). Simultaneously, the employment share of large establishments (≥ 100 employees) declined by roughly 5 percentage points. Thus, either large incumbents reduced their employment, smaller establishments did not grow large (any more) or (very) large establishments predominantly exited from the market. Interestingly, we do not see a rebound of the average size of large establishments in the 2010s (while the share of these establishments rebounds). These findings are consistent with the report of the Monopolies Commission of Germany (2022, p. 23), illustrating that the employment share (and the share of value added) of the 100 largest companies (as opposed to establishments) in Germany declined in the 1990s and 2000s and slightly recovered until 2014.

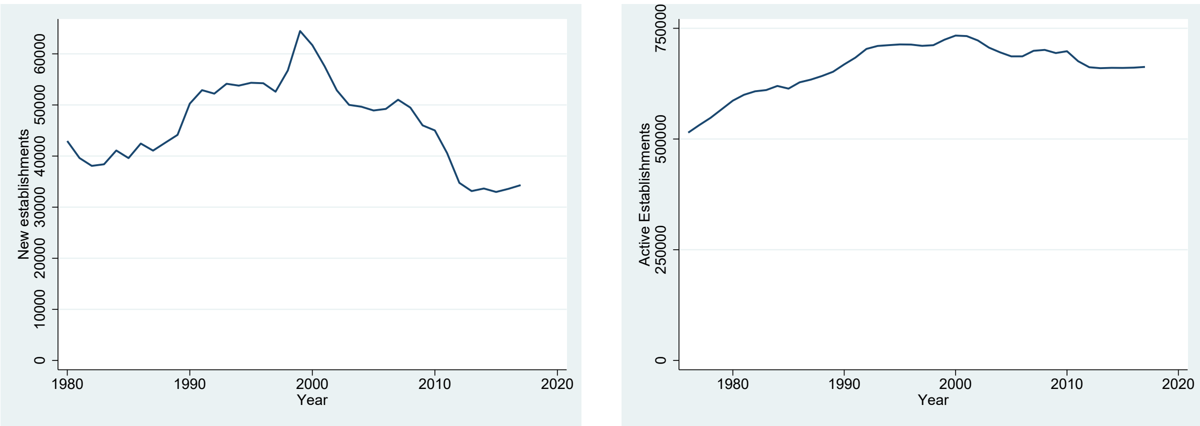
Alongside the decline in the size of large establishments, we present the second contributing factor to the declining average establishment size in Figure 2. As can be seen, the early 1990s experienced a substantial influx of new establishments. While during the 1980s, about 40,000 new establishments entered per year, this number rose to over 50,000 during the 1990s. As new establishments tend to be smaller than incumbents (see, e.g., Fackler et al., 2022), they reduce the average establishment size. At the end of the 2000s, the number of new establishments per year substantially declined again. In most recent years, we observe fewer entries per year than during the 1980s (roughly 32,000). The influx of new establishments resulted in an increasing number of active establishments during the 1990s, as is depicted in the right panel of Figure 2. However, note that the number of active establishments started to increase already in the 1980s.

To quantify the relative importance of the components of the change in average establishment size, we propose a novel decomposition, which isolates size changes of incumbent establishments and changes in the number of entering and leaving establishments as driving forces. The literature mainly applies similar decompositions for variables such as productivity (Foster et al., 2001) or wages (Malchow-Møller et al., 2011).

Define three types of establishments (indexed by j): incumbents C, exits X and entries E, i.e., $j \in \{C, X, E\}$. Consider two periods t and $t+1$. An incumbent establishment is active

Figure 2: Entries and active establishments

(a) Number of establishment entries (b) Number of active establishments



in both periods, t and $t+1$. In contrast, an exiting establishment is active in t but not in $t+1$, while an entry is not active in t but is active in $t+1$. The average establishment size is defined as

$$\bar{n}_t = \frac{1}{M_t} \sum_{i=1}^{M_t} n_{i,t}, \tag{2}$$

where M_t is the number of establishments in t and $n_{i,t}$ is the number of employees of establishment i in t. Define the average establishment size of a specific type $j \in \{C, X, E\}$ as

$$\bar{n}_t^j = \frac{1}{M_t^j} \sum_{i=1}^{M_t^j} n_{i,t}^j \tag{3}$$

and the share of an establishment type in all establishments as

$$\theta_t^j = \frac{M_t^j}{M_t}. \tag{4}$$

Using these definitions, we can decompose a change in the average establishment size between two periods as follows (see appendix for the derivation)

$$
\begin{aligned}
\Delta\bar{n}_{t+1} = & \overbrace{\theta_t^C \Delta\bar{n}_{t+1}^C}^{\substack{\text{Incumbent} \\ \text{size effect}}} + \overbrace{\Delta\theta_{t+1}^C(\bar{n}_t^C - \bar{n}_t)}^{\substack{\text{Incumbent} \\ \text{share effect}}} + \overbrace{\Delta\theta_{t+1}^C \Delta\bar{n}_{t+1}^C}^{\substack{\text{Interaction} \\ \text{term}}} \\
& + \underbrace{\theta_{t+1}^E(\bar{n}_{t+1}^E - \bar{n}_t)}_{\substack{\text{Establishment} \\ \text{entry effect}}} - \underbrace{\theta_t^X(\bar{n}_t^X - \bar{n}_t)}_{\substack{\text{Establishment} \\ \text{exit effect}}} .
\end{aligned}
\tag{5}
$$

The incumbent size effect is the part of the change due to size changes of incumbent establishments (continuing establishments). The incumbent share effect represents the effect of the change in the share of incumbent establishments due to entries and exits. The interaction term has no meaningful interpretation. The establishment entry effect indicates the direct effect of the newly founded establishments on the change of the average establishment size. Lastly, the establishment exit effect represents the direct effect of closures of establishments on the change of average establishment size (please note that the establishment exit effect has a negative sign).

Table 1 displays the decomposition results in absolute and relative terms for five-year periods from 1990 onwards. In addition, it gives the results for the periods 1990-2005 and 2005-2019, where we see the largest changes in the average establishment size. Between 1990 and 1995, the average establishment size decreased by about one employee per establishment. About one-half of this change can be attributed to the decline of the average establishment size of incumbent establishments (incumbent size effect). Hence, incumbent establishments became smaller on average. Also, their employment share declined due to more frequent entry relative to exits, which further decreased the average establishment size (incumbent share effect). The establishment entry effect indicates that the average establishment size would have been about 2.4 employees larger if there had been no entries during this period. Analogously, the establishment exit effect implies that the average establishment size would have been about two employees smaller if there had been no exits. These numbers reflect that many new establishments leave the market shortly after entry. On net, entries and exits contributed slightly less than half of the change in the average establishment size.

Table 1: Average establishment size decomposition

Period		Change	Absolute contribution				
t	$t+1$	$\Delta\bar{n}_{t+1}$	Incum. size	Incum. share	Interact.	Est. entry	Est. exit
1990	1995	-0.965	-0.497	-0.080	0.020	-2.413	-2.005
1995	2000	-0.545	-0.219	-0.069	0.008	-2.135	-1.869
2000	2005	-0.175	-0.243	0.064	-0.008	-1.870	-1.882
2005	2010	0.210	0.277	-0.023	-0.004	-1.791	-1.750
2010	2015	1.532	1.142	0.058	0.039	-1.420	-1.713
1990	2005	-1.682	-0.828	-0.074	0.020	-3.885	-3.085
2005	2019	2.499	1.448	0.180	0.087	-2.201	-2.984

Period		Change	Relative contribution			
t	$t+1$	$\Delta\bar{n}_{t+1}$	Incum. size	Incum. share	Interact.	Net entry
1990	1995	-0.965	0.515	0.082	-0.020	0.422
1995	2000	-0.545	0.401	0.127	-0.015	0.487
2000	2005	-0.175	1.385	-0.364	0.047	-0.069
2005	2010	0.210	1.321	-0.109	-0.017	-0.195
2010	2015	1.532	0.745	0.038	0.025	0.191
1990	2005	-1.682	0.492	0.044	-0.012	0.475
2005	2019	2.499	0.579	0.072	0.035	0.313

Notes: Own calculation based on BHP data (see Ganzer et al., 2021). Incum. size: share of incumbent establishment in period t times change in average establishment size of incumbent establishments. Incum. share: change in the share of incumbent establishments times the period t difference between the average size of incumbent establishments and the average size of all establishments. Interact.: change in the share of incumbent establishments times change in average establishment size of incumbent establishments. Est. entry: share of entries in period $t+1$ times the period $t+1$ difference between the average size of entries and the average size of all establishments. Est. exit: share of exits in period t times the period t difference between the average size of exits and the average size of all establishments.

As can be seen in the third row, the average establishment size declined until 2005, albeit with increasingly smaller changes. However, the relative contributions of the components to the changes remained relatively stable (an exception is the period from 2000 to 2005). Between 2005 and 2010, the average establishment size slowly began to increase again (by roughly 0.2 employees per establishment). The increase in the average establishment size due to the growth of incumbents is larger than the actual increase (0.277 vs. 0.210) since, on net, entries and exits decreased the average establishment size. Hence, the average size would have declined further if incumbents had not increased their workforce. Between 2010 and 2015, the average establishment size increased by more than 1.5 employees per establishment. About 75% of this rise can be attributed to the growth of incumbents. In this period, entries and exits, on net, contributed about 19% to the increase in average establishment size, which is consistent with the stalled entry dynamics visible in Figure 2a.

Overall, between 1990 and 2005, the decline in the size of incumbents and their decreasing share contributed 53.6% to the total decline, while entries and exits accounted for 47.5% (the percentage is over 100 because the interaction term has a positive value).[2] Between 2005 and 2019, the rise in the size of incumbents contributed 57.9% to the total rise and the increasing share of incumbents (due to stalled entry and exit dynamics) contributed 7.2%. On net, entries and exits contributed 31.3% (the latter being due to a rise in the share of exits, which increased the average establishment size since most exits are young and, therefore, small).

5 Establishment size and productivity

Why should we care about the evolution of the average establishment size? In this section, we follow Pagano and Schivardi (2003) and establish a link between aggregate productivity and the average establishment size. The analysis is based on county-level data on labor productivity provided by the German national accounts. We combine these data with county-level measures of average establishment size from the BHP. The estimation equation reads as follows:

$$y_{rt} = \gamma_0 + \gamma_1 S_{rt} + X'_{rt}\gamma + \alpha_r + u_{rt}. \tag{6}$$

Where y_{rt} is labor productivity in county r in period t, S is the corresponding average establishment size, X_{rt} is a vector of control variables, α_r represents county-level fixed effects

[2] Note that with increasingly longer time periods, the contribution of incumbents will tend towards zero as after infinitely many years there will be no incumbents if each establishment has a positive probability of dying. Additionally, many establishments will not be included in the calculation since establishments that are born and die within the same time period are naturally excluded.

and u_{rt} is the error term. The parameter of interest is γ_1, whose estimated coefficient reveals the association between productivity and average establishment size.

We employ two measures of labor productivity: real GDP per capita and real GDP per hour. While GDP per hour is the more accurate measure of labor productivity, the national accounts only provide it from 2000 onward. Therefore, we refer to GDP per capita as our main productivity indicator. With the control variables contained in X_{rt}, we intend to absorb as many confounding factors as possible with the data at hand. In the first set of specifications (rows 1, 2, 4 and 5 in Table 2), we control for the regional employment shares of highly and lowly qualified workers, females, foreigners and routine workers (these shares are based on information provided in the BHP), federal state dummies (in the OLS estimations) and year dummies. In the second set of specifications (rows 3 and 6), we additionally control for employment shares in economic sectors (first digit of the German Classification of Economic Activities 1993) and allow for interaction terms of federal state and year dummies.

The estimation results are displayed in Table 2. The results reveal a positive and statistically significant relationship between productivity and the average establishment size. According to our estimates, the decline in the average establishment size by two employees in the 1990s and 2000s was associated with a substantial decrease of GDP per capita between 2,000 euros (fixed effects estimates) and 3,000 euros (OLS estimates) measured in prices of 2015 (CPI-deflated). Considering GDP per hour as the dependent variable yields the same insights. The decline in the average establishment size by two employees was associated with a decrease of about 0.8 euros (OLS estimates) to 1 euro (FE estimates) per hour of work (measured in CPI-deflated prices of 2015). These sizeable estimates suggest that there may be an association between the decline in the average establishment size and the productivity slowdown in Germany.

However, we cannot interpret these results as revealing causal relations due to reversed causality and potentially omitted variables. For instance, it is conceivable that the average establishment size is influenced by productivity since higher productivity allows establishments

Table 2: Labor productivity and the average establishment size

Dep. Var.:	GDP per capita			GDP per hour		
	OLS	FE	FE	OLS	FE	FE
Avg. size	1528.58***	1167.10***	1226.39***	0.40***	0.52***	0.50***
	(235.42)	(240.21)	(248.90)	(0.06)	(0.13)	(0.13)
Share rout.	-228.11***	-4.50	19.43	-0.03	0.03	0.03
	(70.92)	(23.87)	(36.02)	(0.03)	(0.02)	(0.02)
Share fem.	693.22***	333.66***	335.02***	-0.14*	0.12*	0.18**
	(172.13)	(96.78)	(89.88)	(0.08)	(0.07)	(0.07)
Share for.	145.65	240.87	69.87	0.37***	0.17	0.02
	(174.76)	(197.35)	(212.56)	(0.10)	(0.11)	(0.13)
Share hq.	569.85***	48.78	73.95	0.61***	-0.08	-0.03
	(172.76)	(107.99)	(101.44)	(0.12)	(0.10)	(0.10)
Share lq.	-673.91***	-314.13***	-139.78	-0.17	-0.39***	-0.24**
	(196.14)	(60.26)	(116.13)	(0.11)	(0.08)	(0.10)
Time dummies	✓	✓	✓	✓	✓	✓
State dummies	✓	-	-	✓	-	-
State-time tnt.	-	-	✓	-	-	✓
Sector shares	-	-	✓	-	-	✓
Constant	10037.42	9400.44*	6080.93	40.36***	36.56***	23.53***
	(9587.42)	(5508.80)	(8060.63)	(4.62)	(3.50)	(6.26)
R^2	0.67	0.50	0.55	0.60	0.47	0.52
N	8142	8142	8142	6480	6480	6480

Notes: Robust standard errors clustered on the county level are displayed in parentheses. *** p<0.01, ** p<0.05, * p<0.1 indicate the significance levels. Avg. size: average establishment size, share rout.: share of routine workers, share fem.: share of female workers, share for.: share of foreign workers, share hq.: share of high qualified workers, share lq.: share of low qualified workers and state-time int.: state-time interactions.

to grow larger. Furthermore, with our data at hand, we cannot control for every potential variable that might impact the size-productivity link.

In the appendix, we provide robustness checks based on German labor market regions instead of counties (Tables A.1 and A.2). The labor market regions represent commuting zones (see Breidenbach et al., 2018) and, hence, self-contained local labor markets, making them more appealing to study general equilibrium effects. The results based on labor market regions are very similar to those presented in Table 2.

6 Establishment size distribution and the size wage premium

6.1 Estimation of the regional size wage premium

As previously discussed, we presume that the evolutions of the average establishment size and the size wage premium are interrelated. Therefore, in this section, we study the link between the average establishment size and the size wage premium by exploiting county-level variation. At first, we construct a regional measure of the size wage premium. To do so, we conduct establishment-level estimations for every county and period, regressing a measure of the establishment's wage level (or premium) on its employment size. Since we interpret this relationship as an indicator of monopsony power, the respective coefficient reveals the inverse elasticity of labor supply (Sokolova & Sorensen, 2021). Similar approaches have been conducted by Bachmann, Bayer, et al. (2022), Fakhfakh and Fitzroy (2006), Matsudaira (2014) and Sulis (2011). Note that we face a problem regarding reverse causality since the size of an establishment can potentially be a function of its wages (which would give us estimates of the labor supply elasticity). This approach has been applied by Bodah et al. (2003), Falch (2010) and Staiger et al. (2010). The fact that the coefficients of these two approaches mostly do not align with each other hints towards structural inconsistencies of these two estimation methods (Sokolova & Sorensen, 2021; Manning, 2003). However, a careful assessment of these issues is beyond the scope of this paper.

Specifically, we apply two approaches to measure an establishment's wage level or premium. The first approach uses AKM effects to isolate wage premiums from worker characteristics following Bloom et al. (2018) and Lochner et al. (2020), while the second approach uses average establishment wages and directly controls for other establishment's characteristics.

For the first approach, we take the establishment effect resulting from the estimation of $\psi_{J(i,t)}$ shown in equation 1. It can be interpreted as a wage premium paid by establishment j to all employees independent of their observed and unobserved characteristics (see Bloom et al., 2018; Lochner et al., 2020). To obtain a measure for the regional size wage premium (and, hence, regional monopsony power), we estimate the following equation for each region r and each available AKM interval t separately:[3]

$$\psi_{J(i,t)} = \delta_0 + \delta_1 \log(S_{jt}) + sector_{jt} + v_{jt}. \tag{7}$$

S_{jt} is the average establishment size in terms of full-time employees of establishment j over the period t (e.g., its average size over the period 1985-1992), $sector_{jt}$ represents one-digit sector dummies and v_{jt} is a random error term. δ_1 is the parameter of interest representing the regional size wage premium in period t. We refer to this parameter as swp_{rt} in the following. For the second approach, we use the average wages of full-time employees to measure an establishment's wage level. Here, we cannot directly control for confounding factors related to workers' characteristics since we rely on establishment-level data. However, we account for the most important factors related to the establishment and include the economic sector (one-digit), the shares of engineers, highly qualified and low-qualified employees, the share of routine workers, the shares of females and foreigners and the age of the establishment. It should be stressed that in some region-year cells, the number of establishments is low (on average, there are 305 observations per cell, while the minimum amounts to 31 observations), which reduces the statistical power of our estimates and introduces a sizeable amount of estimation noise. Therefore, we constrain our estimation to the control variables introduced above.

We estimate the following equation, again, for each county r and year t separately

$$\log(\bar{w}_{jt}) = \kappa_0 + \kappa_1 \log(S_{jt}) + X'_{jt}\kappa + \zeta_{jt}. \tag{8}$$

Where $\bar{w}_{jt}$ is the average wage of establishment j in year t, S_{jt} is the size of establishment j in year t in terms of full-time employees, X_{jt} is the vector of control variables described above and ζ_{jt} represents a random error term. Here, the parameter of interest is κ_1, which serves as a

[3]Recall that the AKM effects are available for the intervals 1985-1992, 1993-1999, 1998-2004, 2003-2010 and 2010-2017.

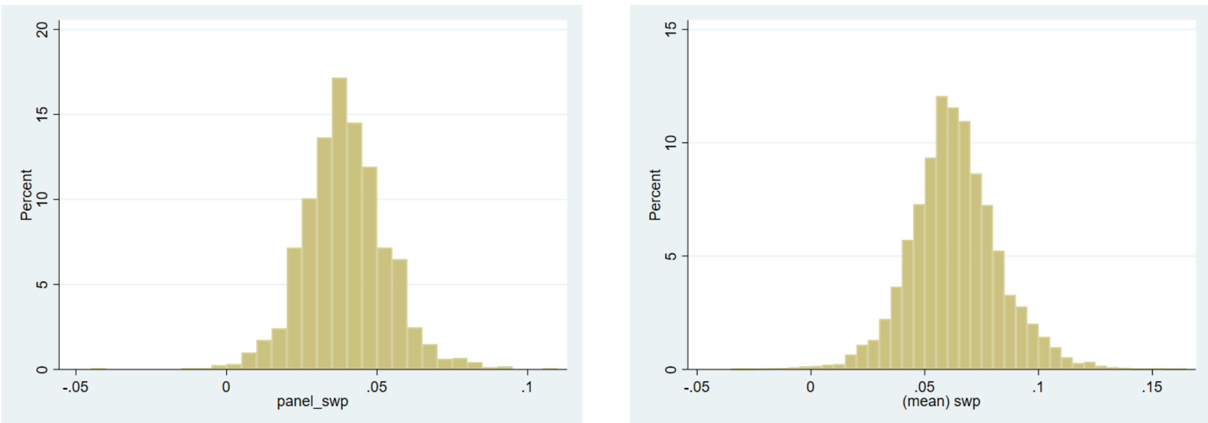
measure for the size wage premium in region r and year t. Hence, it represents a competing (but yearly) measure of swp_{rt}.

For the estimation of both equations (7) and (8), we exclude establishments with less than ten employees, particularly because these small establishments are subject to different institutional regulations that might impact their size-wage-nexus, such as dismissal protection. This sample restriction does, however, not drive our main results. Figure 3 shows histograms of the regional size wage premiums based on AKM and average wages. For both measures, the regional variation in the premium is considerable. Based on the AKM approach, a one percent increase in the size of an establishment is, on average, associated with an increase in the average wage by 0.039% (average over regions and periods). For the average wage approach, this association amounts to 0.063%.

Figure 3: Histogram of regional size wage premiums

(a) AKM-based

(b) Average wage-based



Notes: The bin width is set to 0.05.

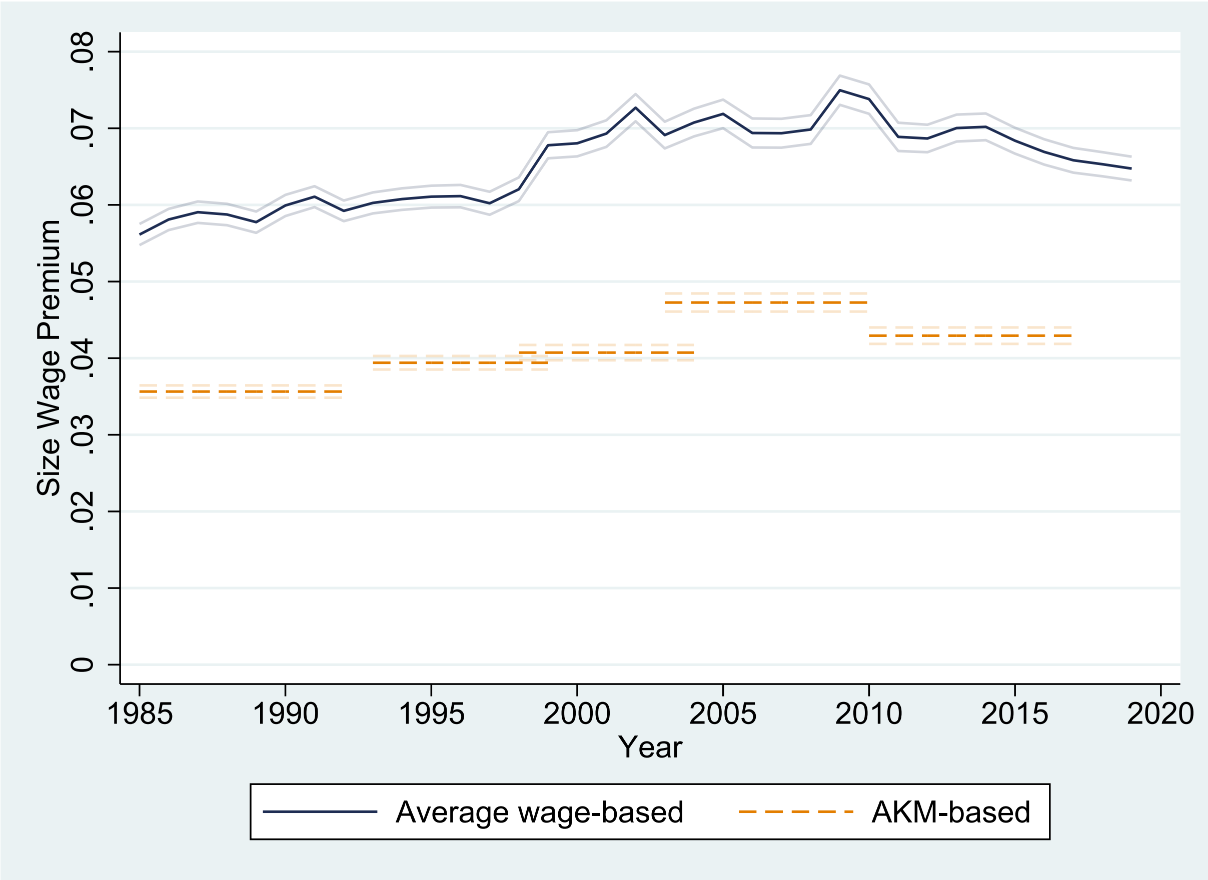
Furthermore, in some rare instances, we estimate negative size wage premiums. Another point worth mentioning is that we estimate the size wage premiums with a measurement error. Hence, estimations with the size wage premiums as explanatory variables will be biased towards zero due to the measurement error problem. Thus, results based on the estimated regional size wage premiums should be interpreted as a lower bound.

6.2 The size wage premium and monopsony power in the labor market

The size wage premiums, estimated in the previous section, serve as our time-varying indicator for monopsony power in regional labor markets. In Figure 4, we depict the evolution of the estimated premiums in West Germany, according to both introduced approaches. Note that the estimations underlying Figure 4 are not based on the regional level but are conducted for each interval/year for West Germany as a whole. The dashed lines correspond to the AKM-based size wage premiums, while the solid line corresponds to the average wage-based size wage premiums. The shaded lines represent the corresponding 95% confidence intervals. As can be seen, there was a statistically significant increase in the size wage premium during the 1990s and (early) 2000s. According to the AKM-based size wage premiums, a one-percentage increase in an establishment's size is, on average, associated with an increase of (worker characteristics adjusted) wages by 0.036% in the period 1985-1992, while this coefficient rises to a maximum of 0.047% in the period 2003-2010 and declines to 0.043% in the period 2010-2017. The rise of the size wage premium was sizeable, amounting to an increase of over 30%, while the subsequent decline amounts to above 8%. According to the average wage-based size wage premiums, a one-percentage increase in an establishment's size is, on average, associated with an increase in average wages by 0.056% in 1985, while this coefficient rose to 0.072% in 2005. In the 2010s, we observe a steady decline to a size wage premium of 0.065% in 2019.

Thus, it can be seen that the evolution of the size wage premiums is independent of the chosen estimation approach. However, the levels between the two measures differ. The size wage premiums, estimated with average establishments wages, are larger than those estimated with the AKM-based approach (by 0.02 to 0.03 percentage points). This may be due to the different worker composition of larger establishments compared to smaller ones, which is accounted for in the AKM-based size wage premiums. Generally, the AKM approach absorbs more worker-related confounding factors, which explains the lower levels of the coefficients. Our

Figure 4: Size wage premium over time in West Germany



Notes: The dashed lines depict the point estimates of the size wage premium based on AKM effects and the solid line depicts the point estimates of the size wage premium based on average wages. Shaded lines represent the corresponding 95% confidence intervals.

results are in line with previous research on the evolution of the size wage premiums in Germany, such as Colonnelli et al. (2018) and Lochner et al. (2020).

Following Bachmann, Bayer, et al. (2022), Green et al. (1996) and Manning (2011, p. 1017-1018), we interpret the size wage premium as a measure of monopsony power in the labor market. In the modern view, monopsony power in the labor market is reflected by an upward-sloping labor supply curve to the firm. Since the size wage premium measures the slope of the labor supply curve, it can be regarded as the inverse labor supply elasticity a firm is confronted with. Put differently, it reflects the wage increase a firm would have to pay if it wants to increase its labor input. Analogously, the labor supply elasticity measures the amount of labor lost if the firm wants to decrease its labor compensation. As a result, employers react to rising monopsony power by reducing their workforce, thereby maximizing their monopsony rents. As shown in Figure 1 and Table 1, this is exactly what we observe in the period of rising monopsony power (as measured by the size wage premiums) and the opposite of what we observe in the period of declining monopsony power.

However, as discussed in section 2, there are many competing explanations for the existence of a size wage premium, such as the selection of high-skilled workers into large firms (which we account for in our approach), efficiency wages, compensating differentials and rent sharing (for an overview, see Colonnelli et al., 2018; Green et al., 1996). To further strengthen our argument that the size wage premium reflects monopsony power in the labor market, we compare the evolution of the size wage premium with other measures for monopsony power. According to Manning (2003, pp. 44-49) and Hirsch et al. (2022), an alternative indicator for monopsony power is the share of hires from non-employment. In a competitive labor market, employers compete directly for workers, which leads to poaching and, hence, to many job-to-job transitions. With decreasing competition (i.e., increasing monopsony power) in the labor market, poaching will decline as employees are increasingly tied to their employers. Thus, employers will have to rely more on hires from non-employment to adjust their labor input (see Hirsch et al., 2022).

Using the public release data of the Administrative Wage and Labor Market Flow Panel (see Stüber & Seth, 2019), we can compute the share of hires from non-employment for West Germany on a yearly basis from 1976 to 2014. We depict this share in Figure 5 and additionally include a smoothed trend based on the HP-filter (smoothing parameter $= 100$ due to the yearly frequency) to illustrate the general evolution. Figure 5 reveals a rising share of hires from non-employment in the 1990s and early 2000s and a declining share beginning in the late 2000s. Hence, according to this measure, the labor market became less competitive during the 1990s and early 2000s and more competitive afterwards. Thus, the overall picture is consistent with our interpretation of the size wage premium: monopsony power rose during the 1990s and early 2000s and declined afterwards.

Figure 5: Rate of hires from non-employment

Notes: This figure is based the public release data of the Administrative Wage and Labor Market Flow Panel (see Stüber & Seth, 2019), own calculations.

Further, the evolution of our two measures of monopsony power, illustrated in Figure 4 and Figure 5, is consistent with findings in the literature. For instance, Hirsch et al. (2018) find that monopsony power in West Germany substantially increased during the 1990s, partly recovered around the turn of the millennium and declined after 2005. In their framework, the imperfect elasticity of labor supply to the firm stems from search frictions, which are stronger in economic downturns. As is standard in this type of framework, the equilibrium search model of Burdett and Mortensen (1998) is used to derive an estimation strategy that aims to measure the labor supply elasticity to the firm with the wage elasticities of recruitment and separation rates (for a detailed description, see Manning, 2011). Bachmann, Demir, and Frings (2022) applied the same approach but differentiated between workers with three different routine task intensities.

Their results on the evolution of monopsony power align with Hirsch et al. (2018) and our results. Furthermore, even though there are differences in the level the three worker groups are exposed to monopsony, the evolution is the same for all workers. Hence, the described changes in monopsony power are not due to shifts in the worker or task composition in the West German economy.

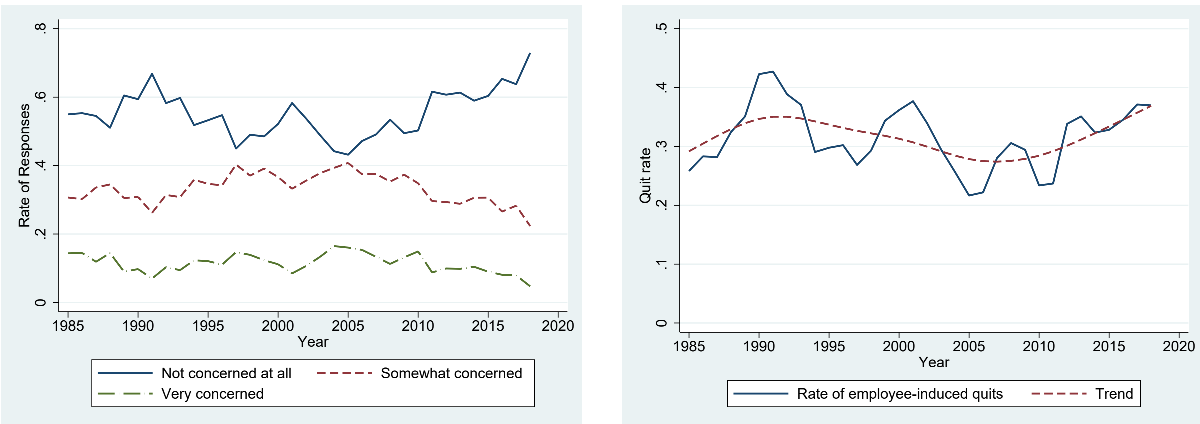
Why did monopsony power rise during the 1990s and decline in the 2010s in West Germany? Both Hirsch et al. (2018) and Bachmann, Demir, and Frings (2022) conclude that monopsony power varies counter-cyclically, i.e. that it is higher in times of high unemployment rates. Further, Bachmann, Demir, and Frings (2022) argue that long term developments, such as technological change or labor market polarization, do not play major roles in driving monopsony power. Another reason for the rise in monopsony power during the 1990s may be the increasing international trade volume. With the fall of the iron curtain, employers may have been able to credibly threaten to shift operations to Eastern Europe (see, e.g., Brändle, 2015), thereby redistributing labor market power. However, there is no plausible explanation for why such threats should have lost their credibility in the 2010s. Furthermore, Moser et al. (2015) demonstrate that establishments in Germany rather increased employment if they engaged in offshoring as productivity gains exceeded the direct employment losses from offshoring.

Generally, the rise in monopsony power falls in a time when Germany was called the "Sick Man of Europe", while the decrease in monopsony power can be associated with the revival of the German economy in the aftermath of the Great Recession (see Dustmann et al., 2014). Why should these developments be related to monopsony power? The idea is simple: in economically challenging times, workers' outside options may be limited such that they would rather favor job security than potential wage gains through job changes. Consequently, employers would gain monopsony power (see Hirsch et al., 2018).

To gain further insights into the evolution of outside options, we use another data source, the German Socio-Economic Panel (see Goebel et al., 2019), which provides us with survey-based information on concerns about job security and employee quitting behavior. Specifically, we investigate the question "How concerned are you about the following issues? If you are employed: Your job security" (Figure 6a) and the rate of (employee-induced) quits relative to all other causes of job destruction (Figure 6b). As can be seen, an increasing share of employed individuals answered to be at least somewhat concerned in the 1990s and 2000s, while a decreasing share of employed individuals answered to be not concerned at all. Simultaneously, quit rates substantially declined during the 1990s and early 2000s and increased in the late 2000s and 2010s.[4] Both pieces of evidence give us a consistent picture which is in line with our previous argumentation: employees became increasingly concerned about their job security and these concerns seem to have translated into a more conservative and job-preserving quitting behavior. This observation may indicate a declining number of outside options for the current job. Consequently, the individuals were more and more tied to their employers, which increased the monopsony power of the firm. An exciting avenue for further research would be to clarify the causal relationship underlying these developments. This question is, however, beyond the scope of our paper.

Figure 6: Job security and quit rate

(a) Concerns about job security (b) Quit rate



[4]Note that the evolution of the quit rates is roughly the inverse of the evolution of the share of hires from non-employment, shown in Figure 5. This observation makes sense since fewer employee-induced quits imply less scope for poaching and, therefore, more hires from non-employment.

6.3 The size wage premium and average establishment size

In this section, we empirically investigate the link between the average establishment size and the size wage premium. To do so, we estimate the following equation on the county-level

$$S_{rt} = \lambda_0 + \lambda_1 swp_{rt} + X'_{rt}\lambda + \eta_r + \varepsilon_{rt}, \tag{9}$$

where S_{rt} is the average establishment size in region r at time period t, swp_{rt} is the size wage premium from the estimation of equation (7) and equation (8), respectively, X_{rt} is a vector of control variables, η_r is the regional fixed effect, ε_{rt} is the error term and the λs are the parameters to be estimated. λ_1 indicates the link between average establishment size and size wage premium and, therefore, represents the parameter of interest.

For our estimations, we apply both measures of the size wage premium we introduced above. Note that in the AKM-based regressions, the observational unit is a county in a time interval (instead of a specific year). Therefore, we measure the average establishment size in the last year of the period and all control variables in the first year of the period. We estimate the parameters of the model using both simple OLS and fixed effects estimators for both measures. As control variables, we include the regional employment shares of high-qualified, low-qualified, females, foreigners and routine workers (these shares are based on information provided in the BHP), federal state dummies (in the OLS estimations) and period dummies or year dummies. In our second fixed effects specification, we additionally control for employment shares in economic sectors (first digit of the German Classification of Economic Activities 1993) and allow for interaction terms of federal state dummies and year dummies. In total, we conduct six estimations, three for each measure of the size wage premium.

Table 3 indicates that in each specification, the parameter of the size wage premium (first row) is statically significant and has the expected negative sign. Thus, with our estimations, we can document a robust negative link between the size wage premium in a county and its respective average establishment size. In terms of the economic magnitude of our estimates,

Table 3: Average establishment size and the size wage premium

Dependent variable: average establishment size

	AKM-based			Average wage-based		
	OLS	FE	FE	OLS	FE	FE
swp_{rt}	-221.44***	-67.72**	-56.49**	-159.27***	-42.40***	-36.83***
	(72.53)	(26.33)	(23.27)	(57.83)	(13.05)	(10.78)
Share rout.	24.72	-17.30**	-19.41**	5.80	-6.18	-7.12
	(25.16)	(8.55)	(9.29)	(19.69)	(5.80)	(7.82)
Share fem.	-220.02***	-46.09***	-16.45	-237.23***	-100.38***	-67.92***
	(50.81)	(14.03)	(15.92)	(56.27)	(16.19)	(15.92)
Share for.	-2.68	65.24*	49.42	-19.54	40.63***	29.48**
	(42.87)	(37.75)	(34.43)	(29.99)	(12.89)	(13.41)
Share hq.	239.37***	-0.08	-21.04	208.56***	7.68	-19.87
	(64.26)	(19.82)	(20.60)	(57.89)	(29.85)	(38.93)
Share lq.	34.71	-3.53	-10.49	56.50	29.87*	16.91
	(43.89)	(21.77)	(22.65)	(42.63)	(15.97)	(19.46)
Time dummies	✓	✓	✓	✓	✓	✓
State dummies	✓	-	-	✓	-	-
State-time int.	-	-	✓	-	-	✓
Sector shares	-	-	✓	-	-	✓
Constant	86.21***	82.16***	79.77***	98.15***	79.30***	82.56***
	(12.12)	(6.99)	(12.05)	(14.99)	(4.78)	(11.61)
R^2	0.23	0.18	0.25	0.22	0.12	0.28
N	1620	1620	1620	11340	11340	11340

Notes: Robust standard errors clustered on the county level are displayed in parentheses. *** p<0.01, ** p<0.05, * p<0.1 indicate the significance levels. swp_{rt}: size wage premium, share rout.: share of routine workers, share fem.: share of female workers, share for.: share of foreign workers, share hq.: share of high qualified workers, share lq.: share of low qualified workers and state-time int.: state-time interactions.

we associate the size of the coefficients with the observed change in the size wage premium between the early 1990s and the mid-2000s (i.e., $\lambda_1 \times \Delta\text{swp}$). Starting with the AKM-based OLS estimations, the observed change in the size wage premium between the intervals 1985-1992 and 2003-2010 ($\Delta\text{swp} = 0.012$) is associated with a decline in the average establishment size by 2.68 employees per establishment. The respective fixed effects estimates are associated with a decline of 0.82 and 0.68 employees per establishment, respectively. The actual decline in size between 1990 and 2005 was about 1.7 employees per establishment. Hence, even with the smaller estimates from the fixed effects regressions, we can explain about 40% of the decline in average establishment size with the rise in the size wage premium. Furthermore, these estimates imply an increase in the average establishment size in the 2010s by 0.99 (OLS), 0.30 (FE) and 0.25 (FE) employees, respectively. The actual increase from 2010 to 2017 amounts to 1.89 employees per establishment. Thus, the evolution of the size wage premium is also related to the rebound in average establishment size, although the relationship is less pronounced here.

In the average wage-based estimations, the OLS estimate implies that the change in the size wage premium from 1990 to 2005 is associated with a decline of the average establishment size by 1.90 employees per establishment. The respective fixed effects estimates are associated with a decline of 0.51 and 0.44 employees per establishment, respectively. Even though the associated decline in the average establishment size is markedly smaller in this case, we can still explain at least one-fourth of the decline in the average establishment size. Furthermore, the evolution of the size wage premium is associated with a rebound of the average establishment size in the 2010s by 1.44 (OLS), 0.38 (FE) and 0.33 (FE) employees per establishment. These results are consistent with Bachmann, Bayer, et al. (2022), who show that differences in the size wage premium in West and East Germany are associated with differences in the establishment size distribution.

As shown in Table 1, establishment entries are an important contributor to the change in the average establishment size. Therefore, we investigate the link between the prevalence of establishment entries and the size wage premium in the following, presuming a positive relationship. The economic rationale is that in a region with a large size wage premium, it may be advantageous to enter the market as a new and small establishment since higher monopsony power (indicated by higher size wage premiums) promises higher rents.[5]

To provide evidence on this link, we repeat the estimations from above and replace our dependent variable with the number of new establishments per employee in a region. More precisely, in the AKM-based regressions, the dependent variable is the sum of new establishments

[5]Another view is that new establishments can only pay lower wages because these establishments do not (yet) achieve economies of scale and would, therefore, not be profitable if they paid the same wages as incumbents. Therefore, a steep size-wage relation may encourage establishment entries by increasing the probability of survival.

in a region r over an AKM interval divided by the number of employees in this region in the last year of the period. In the average wage-based regressions, the dependent variable is the sum of entries in a region and year divided by the number of employees in this region and year. Independent variables and specifications remain the same as in Table 3 .

As expected, Table 4 indicates a positive and statistically significant relationship between the prevalence of new establishments in a region and the regional size wage premium. The coefficient of the size premium based on the AKM approach amounts to 0.308 in the OLS

Table 4: Entries per employee and the size wage premium

Dependent variable: entries per employee

	AKM based			average wage based		
	OLS	FE	FE	OLS	FE	FE
swp_{rt}	0.308***	0.071*	0.070**	0.038***	0.014***	0.012***
	(0.072)	(0.040)	(0.035)	(0.005)	(0.002)	(0.002)
Share rout.	0.008	0.007	0.009	0.003*	-0.000	-0.000
	(0.023)	(0.011)	(0.012)	(0.002)	(0.001)	(0.001)
Share fem.	0.210***	0.068**	0.057**	0.027***	0.014***	0.012***
	(0.032)	(0.028)	(0.028)	(0.004)	(0.003)	(0.003)
Share for.	0.167***	-0.095***	-0.083**	0.021***	-0.008***	-0.011***
	(0.046)	(0.032)	(0.032)	(0.004)	(0.003)	(0.003)
Share hq.	-0.200***	-0.028	-0.037	-0.018***	-0.001	-0.004
	(0.055)	(0.029)	(0.033)	(0.004)	(0.003)	(0.003)
Share lq.	-0.205***	-0.096***	-0.063***	-0.027***	-0.015***	-0.009***
	(0.042)	(0.025)	(0.024)	(0.005)	(0.003)	(0.003)
Time dummies	✓	✓	✓	✓	✓	✓
State dummies	✓	-	-	✓	-	-
State-time int.	-	-	✓	-	-	✓
Sector shares	-	-	✓	-	-	✓
Constant	0.071***	0.082***	0.035	0.006***	0.008***	0.003
	(0.019)	(0.011)	(0.022)	(0.002)	(0.001)	(0.002)
R^2	0.38	0.49	0.59	0.46	0.57	0.63
N	1620	1620	1620	11340	11340	11340

Notes: Robust standard errors clustered on the county level are displayed in parentheses. *** $p<0.01$, ** $p<0.05$, * $p<0.1$ indicate the significance levels. swp_{rt}: size wage premium, share rout.: share of routine workers, share fem.: share of female workers, share for.: share of foreign workers, share hq.: share of high qualified workers, share lq.: share of low qualified workers and state-time int.: state-time interactions.

estimation and 0.070 and 0.071 in the fixed effects estimations. Relative to the average entries per employee over all regions and AKM intervals, these coefficients imply an increase by 1.2% (FE estimates) to 5.2% (OLS estimate) if we multiply the respective coefficient with the change in the size wage premium between the periods 1985-1992 and 2003-2010. On the other hand, our estimates imply a decrease in entries per employee by 0.44% to 1.9% between 2003-2010 and 2010-2017.

The results based on the average wages approach imply increases of new establishments per employee by 1.6%, 2.0% (FE estimates) and 5.2% (OLS estimate) between 1990 and 2005. From 2010 to 2019, the estimates imply a decrease in entries per employee by 1.2% to 4.0%. Hence, the implications of the results based on both approaches are similar and confirm that higher size wage premiums (and, thus, monopsony power) played a role in fostering new business formation in the 1990s and early 2000s, while the decline in the size wage premium in the 2010s is associated with a decline in business formation. Note that the smaller coefficients with average wage-based size wage premiums are due to the smaller number of entries in a year compared to an AKM interval.

To summarize, in this section, we established a statistically and economically significant negative relationship between the average establishment size of a region and the regional size wage premium as well as a positive relationship between the prevalence of new establishments in a region and the regional size wage premium. Therefore, our results suggest that the decline in the average establishment size in the 1990s and 2000s is associated with the rise in the size wage premium, while the rebound of the average establishment size in the 2010s is associated with the decline of the size wage premium. Further, we corroborate our findings from Table 1 by documenting that newly entering establishments have also impacted the variation in the average establishment size. Our results are robust to considering local labor market regions (based on commuting zones) instead of counties, which we present in the appendix (A.2).

7 Model

We propose a model with heterogeneous firms that is able to replicate our empirical findings. The model helps us organize thoughts on potential channels and provide the reader with a systematic analysis of the mechanics behind the fall and rebound of the average establishment size. Furthermore, due to a lack of a credible instrument, we cannot interpret our empirical results causally. However, our model allows us to corroborate our empirical results because this model can replicate the empirical findings with a standard calibration.

The labor market in this model is characterized by monopsonistic competition, which is reflected by an upward-sloping labor supply curve and, hence, a positive firm size wage relation. Monopsonistic competition in the labor market results from the households' love-for-variety for employers. The intuition is that households prefer to choose a suitable job for each household member. With an increasing number of employers, the chances of finding a suitable job for each household member increase. Hence, the household's disutility of working decreases. These relations imply that employers are no perfect substitutes from the households' perspective. Furthermore, the product market in our model is characterized by monopolistic competition and firms can enter the market endogenously and exit the market with an exogenous probability (the exit probability is the same for all existing firms).

7.1 Household

The household consumes a bundle of final goods C_t, which consists of many goods $q(\omega)$, $\omega \in \Omega_t$. Where Ω_t is the set of available goods in a given period, a firm can be indexed by ω as each produces only one good. The household maximizes expected lifetime utility subject to its budget constraint:[6]

[6]Originally, the household's budget constraint is expressed in terms of money. As with the firm's equations in the next sections, the budget constraint is normalized by the price of a consumption bundle P_t.

$$
\max_{C_t, N_t, b_{t+1}, x_{t+1}} E_0 \sum_{t=0}^{\infty} \beta^t \left[\frac{C_t}{1-\gamma}^{1-\gamma} - \bar{\varphi}^{\frac{1}{\varphi}} \frac{N_t}{1+\frac{1}{\varphi}}^{1+\frac{1}{\varphi}} \right]
$$

$$
s.t. \tag{10}
$$

$$
b_{t+1} + \tilde{v}_t (M_t + M_t^e) x_{t+1} + C_t = (1+r_t) b_t + \tilde{v}_t M_t x_t + \tilde{d}_t M_t^o x_t + W_t N_t.
$$

Where $\beta \in (0,1)$ is the discount factor, γ is the elasticity of intertemporal substitution, φ is the Frisch elasticity of labor supply and W_t is the wage index such that $W_t N_t$ is the real labor income. The household finances the firms by buying the shares x_t. More precisely, in period t, the household can purchase shares x_{t+1} of existing firms M_t and new firms M_t^e weighted with the average value of a firm $\tilde{v}_t$. We define M_t^o as operating firms since a fraction of the firms will not operate. Like in Ghironi and Melitz (2005) and Colciago and Silvestrini (2022), the household gets dividends from operating firms $\tilde{d}_t M_t^o$ as a return on the bought share, where $\tilde{d}_t$ is the average dividend of an operating firm. The household continues to finance all existing firms as the household does not know, which firms will exit the market. In addition, the household makes a saving decision over one-period bonds b_t. The return on bonds is the gross interest rate $(1 + r_t)$.

Intertemporal utility maximization yields the Euler equation for bonds

$$
C_t^{-\gamma} = E_t \beta (1 + r_{t+1}) C_{t+1}^{-\gamma}, \tag{11}
$$

the labor leisure condition

$$
W_t = \bar{\varphi}^{\frac{1}{\varphi}} \frac{N_t^{\frac{1}{\varphi}}}{C_t^{-\gamma}}, \tag{12}
$$

and the Euler equation for shares

$$
\tilde{v}_t = E_t \beta (1-\delta)((1 - G(z_{t+1}^c)) \tilde{d}_{t+1} + \tilde{v}_{t+1}) \left(\frac{C_{t+1}}{C_t} \right)^{-\gamma}, \tag{13}
$$

where $(1-G(z^c_{t+1}))$ denotes the fraction of operating firms from the set of all firms and δ denotes the exogenous probability of a market exit, which is the same for all existing firms.

As in Berger et al. (2022) and Wolf (2020), the overall disutility index from working N_t is a result of the household's labor supply to many different firms

$$N_t = \left(\int_{\omega \in \Omega_t} n_t(\omega)^{\frac{\eta+1}{\eta}} d\omega \right)^{\frac{\eta}{\eta+1}}, \tag{14}$$

where n_t is the labor supply to firm ω and η is the elasticity of substitution of labor supply to different firms. Given the decision on N_t, the household decides how much labor it supplies to each firm $n_t(\omega)$. Thus, the household maximizes its labor income subject to the preference of having multiple employers for the household members

$$\begin{aligned} \max_{n_t(\omega)} \int_{\omega \in \Omega_t} n_t(\omega) w_t(\omega) d\omega \quad s.t. \quad N_t^{\frac{\eta+1}{\eta}} &= \int_{\omega \in \Omega_t} n_t(\omega)^{\frac{\eta+1}{\eta}} d\omega \\ \Rightarrow w_t(\omega) &= \left(\frac{n_t(\omega)}{N_t} \right)^{\frac{1}{\eta}} W_t \\ \Leftrightarrow n_t(\omega) &= \left(\frac{w_t(\omega)}{W_t} \right)^{\eta} N_t. \end{aligned} \tag{15}$$

The solution to this problem implies that the household has to be compensated with ever-increasing wages if it is supposed to supply an increasing portion of its labor to a firm. Put differently, a firm has to pay higher wages to attract more employees. This describes the theoretical mechanism behind the positive relation between wages and employment, where $\frac{1}{\eta}$ corresponds to the empirically estimated size wage premium. The implication is that the firms face an upward-sloping labor supply curve, which they internalize in their labor demand decision (shown below). Plugging this result into the definition of the labor supply index gives the wage index

$$W_t = \left(\int_{\omega \in \Omega_t} w_t(\omega)^{1+\eta} d\omega \right)^{\frac{1}{1+\eta}}. \tag{16}$$

The household combines the different final goods into one bundle with a CES utility function

$$C_t = \left(\int_{\omega \in \Omega_t} q_t(\omega)^{\frac{\sigma-1}{\sigma}} d\omega \right)^{\frac{\sigma}{\sigma-1}}, \tag{17}$$

with $\sigma > 1$. Thus, the household's demand for one intermediate good is

$$q_t^H(\omega) = \left(\frac{p_t(\omega)}{P_t} \right)^{-\sigma} C_t = \rho_t(\omega)^{-\sigma} C_t, \tag{18}$$

with $\rho_t(\omega) = \frac{p_t(\omega)}{P_t}$ and

$$P_t = \left(\int_{\omega \in \Omega_t} p_t(\omega)^{1-\sigma} d\omega \right)^{\frac{1}{1-\sigma}}, \tag{19}$$

which is the price index for one consumption bundle.

7.2 Firms

Firms have to pay entry costs f_e to enter and fixed costs f_x to be able to produce. They pay these costs in the form of goods specifically produced for this purpose. These goods are produced in a perfectly competitive sector and the inputs are the good varieties $q(\omega)$. The production function is

$$X_t = \frac{1}{M_t^{o\frac{1}{\sigma-1}}} \left(\int_{\omega \in \Omega_t} q_t(\omega)^{\frac{\sigma-1}{\sigma}} d\omega \right)^{\frac{\sigma}{\sigma-1}}. \tag{20}$$

This setup is similar to Karabarbounis and Neiman's (2014) differentiation between a consumption good sector and an investment good sector. Here the main difference is that while the household itself has a variety effect in its aggregation of the final goods to one bundle, the firm producing the corporate good $q_t^X(\omega)$ does not. This normalization of the aggregator is also used by Felbermayr et al. (2018), for instance. The demand for one corporate good is

$$q_t^X(\omega) = M_t^{o\frac{1}{\sigma-1}} \rho(\omega)^{-\sigma} X_t. \tag{21}$$

Using equation (19), the price for one corporate good can be written as

$$P_t^X = M_t^{o\frac{1}{\sigma-1}} P_t = \xi_t P_t, \tag{22}$$

such that $\xi_t = M_t^{o\frac{1}{\sigma-1}}$ is the relative price of a corporate good to one consumption bundle. Overall demand for one variety is

$$q_t(\omega) = \rho_t(\omega)^{-\sigma}(C_t + M_t^{o\frac{1}{\sigma-1}} X_t) = \rho_t(\omega)^{-\sigma}(C_t + \xi_t X_t) = \rho_t(\omega)^{-\sigma} D_t, \tag{23}$$

where $D_t = C_t + \xi_t X_t$ is a measure for overall demand. Given that M_t^o firms pay the fixed costs f_x and M_t^e firms pay the entry costs f_e, the market clearing condition for the corporate good is

$$X_t = f_e M_t^e + f_x M_t^o. \tag{24}$$

Each firm produces with the aggregate productivity Z_t and the firm-specific productivity z. The firm-specific productivity level is constant over time and is drawn from the distribution $G(z)$ upon the firm's entry. We assume that $G(z)$ follows a Pareto distribution

$$G(z) = 1 - \left(\frac{z_m}{z}\right)^k. \tag{25}$$

A firm with productivity z produces

$$q_t(z) = n_t(z) Z_t z \tag{26}$$

units of output. Since the firm-specific productivity is the only ex-ante difference between the firms, it is now used to index a specific firm. The firms are in monopolistic competition on the goods market and in monopsonistic competition on the labor market. Each firm maximizes profits by choosing production $q_t(z)$, taking into account the demand function for its good and the wage employment relationship. Real profits are

$$\pi_t(z) = q_t(z)\rho_t(z) - w_t(z)n_t(z) - \xi_t f_x. \tag{27}$$

Firms internalize the size wage relation in their profit maximization problem. Thus, maximizing equation (27), taking into account the size wage relationship (15) and the individual firms demand function (23), gives us

$$\rho_t(z) = \frac{\sigma}{\sigma-1}\frac{\eta+1}{\eta}\frac{w_t(z)}{Z_t z}. \tag{28}$$

Hence, the real price a firm is charging for its good is a markup over the costs of an effective unit of labor firstly because of the monopolistic competition on the goods market like in Melitz (2003) and secondly because of the monopsonistic competition on the labor market like in Jha and Rodriguez-Lopez (2021). Using the demand function for one intermediate good, a firm's production function and the labor supply equation for one firm, we can get an expression for the wage

$$w_t(z) = \left(\frac{\sigma(\eta+1)}{(\sigma-1)\eta}\right)^{-\frac{\sigma}{\eta+\sigma}} (Z_t z)^{\frac{\sigma-1}{\eta+\sigma}} \left(\frac{D_t}{N_t W_t^{-\eta}}\right)^{\frac{1}{\eta+\sigma}}. \tag{29}$$

Define operating profits $\pi_t^o(z)$ as firms' profits ignoring the fixed costs $\xi_t f_x$. Hence, operating profits are those after the firm pays the fixed costs to be able to produce. The operating profits of an active firm are

$$\pi_t^o(z) = \left[\frac{((\sigma-1)\eta)^{\eta(\sigma-1)} D_t^{\eta+1} (N_t W_t^{-\eta})^{\sigma-1}}{(\sigma(\eta+1))^{\sigma(\eta+1)}}\right]^{\frac{1}{\sigma+\eta}} (\sigma+\eta)(Z_t z)^{\frac{(\eta+1)(\sigma-1)}{\eta+\sigma}}. \tag{30}$$

7.3 Entry and exit

At the beginning of each period, the existing firms M_t decide if they become active. A firm will become active if it does not make a loss, i.e., the fixed costs of production $\xi_t f_x$ do not exceed the operating profits π_t^o. Given that operating profits of individual firms are increasing in their firm-specific productivity z, there will be a cut-off firm-specific productivity level z_t^c determining if firms become active. The least productive active firm is making zero profits, which gives us the zero-cut-off-profit condition

$$\pi_t^o(z_t^c) = \xi_t f_x. \tag{31}$$

This equation gives us the productivity threshold to become active z_t^c. Since profits vary with, e.g., aggregate productivity Z_t, previously active firms can become inactive as the productivity threshold z_t^c varies. However, these firms do not exit the market altogether but wait for improving economic conditions to start operations again. Like Ascari et al. (2021), we call such firms idle.

There is free entry into the market. It follows that entry occurs until the value of entering the market

$$\tilde{v}_t = E_t \sum_{s=t+1}^{\infty} (\beta(1-\delta))^{s-t}(1 - G(z_s^c))\left(\frac{C_s}{C_t}\right)^{-\gamma} \tilde{d}_s \tag{32}$$

is equal to the entry costs $\xi_t f_e$. Thus, the free entry condition implies

$$\tilde{v}_t = \xi_t f_e. \tag{33}$$

In each period, a fraction δ of firms exit the market exogenously. Hence, the number of firms in period t consists of the surviving firms in $t-1$ and surviving entries in $t-1$. We can write the law of motion for the number of firms as

$$M_t = (1-\delta)(M_{t-1} + M_{t-1}^E). \tag{34}$$

The operating firms are a fraction of the existing firms. This fraction is determined by the productivity threshold to become active z_t^c and the productivity distribution

$$M_t^o = (1 - G(z_t^c))M_t. \tag{35}$$

7.4 Equilibrium

Like Jha and Rodriguez-Lopez (2021), we can define the average productivity in the economy as

$$\tilde{z}_t = \left(\int_{z_t^c}^{\infty} z^{\frac{(\eta+1)(\sigma-1)}{\eta+\sigma}} \frac{g(z)}{(1 - G(z_t^c))} dz \right)^{\frac{(\eta+\sigma)}{(\eta+1)(\sigma-1)}}. \tag{36}$$

Intuitively, this expression is the average over the firm-specific productivities. Define a new parameter to make notation easier $\kappa = \frac{(\eta+1)(\sigma-1)}{\eta+\sigma}$, such that

$$\tilde{z}_t = \left(\frac{k}{k - \kappa} \right)^{\frac{1}{\kappa}} z_t^c. \tag{37}$$

The average productivity is an increasing function of the productivity cut-off. Like in Melitz (2003), the exit from the production of relatively unproductive firms increases the average productivity. We can use the aggregate price index to get one more equilibrium condition

$$P_t = \left(\int_{z_t^c}^{\infty} p_t(z)^{1-\sigma} M_t^o \frac{g(z)}{(1 - G(z_t^c))} dz \right)^{\frac{1}{1-\sigma}}. \tag{38}$$

Dividing the equation by P_t and using straightforward algebra yields

$$1 = M_t^{o\,\frac{1}{1-\sigma}} \rho_t(\tilde{z}_t). \tag{39}$$

The dividends paid out to the household by the average firm is equal to the profits of that firm

$$\tilde{d}_t = \pi_t(\tilde{z}_t) = \left[\frac{((\sigma-1)\eta)^{\eta(\sigma-1)} D_t^{\eta+1} (N_t W_t^{-\eta})^{\sigma-1}}{(\sigma(\eta+1))^{\sigma(\eta+1)}}\right]^{\frac{1}{\sigma+\eta}} \times (\sigma+\eta)(Z_t \tilde{z}_t)^{\frac{(\eta+1)(\sigma-1)}{\eta+\sigma}} - \xi_t f_x. \tag{40}$$

Using that bonds are on zero net supply and that the household buys all the shares of the firms in equilibrium, gives us

$$C_t + \xi_t f_e M_t^E = W_t N_t + \tilde{d}_t M_t^o. \tag{41}$$

A second aggregate condition comes from the fact that aggregate demand consists of consumption and the firm's demand for producing the corporate good

$$D_t = C_t + \xi_t (M_t^E f_e + M_t^o f_x). \tag{42}$$

To see how the average employment reacts to changes in the size wage premium, we have to derive an explicit function for the average firm size. In a first step, we derive aggregate employment. Because of the variety effect included in (14), aggregate employment is not equal to N_t. The employment of all operating firms is

$$\bar{N}_t = \int_{z_t^c}^{\infty} n_t(z) M_t^o \frac{g(z)}{(1-G(z_t^c))} dz. \tag{43}$$

Using the labor supply equation of a firm as well as the wage equation of a firm and using some algebra gives us

$$\bar{N}_t = M_t^o \left[\left(\frac{(\sigma-1)\eta}{\sigma(\eta+1)}\right)^{\eta\sigma} \frac{Z_t^{\eta(\sigma-1)} D_t^{\eta} N_t^{\sigma}}{W_t^{\eta\sigma}}\right]^{\frac{1}{\eta+\sigma}} \frac{k}{k-\nu} z_t^{c\nu}, \tag{44}$$

with $\nu = \frac{\eta(\sigma-1)}{\eta+\sigma}$. Thus, average employment per operating firm is

$$\bar{n}_t = \frac{\bar{N}_t}{M_t^o}. \tag{45}$$

7.5 Calibration

We calibrate the model to a quarterly frequency. To achieve this, we set the households discount factor β to 0.99, a standard value on this frequency. We also choose standard values for the elasticity of intertemporal substitution γ by setting it to 1 (e.g., used by Cacciatore et al., 2021) and the Frisch elasticity ϕ, which we set to 0.5 (e.g., used by Eggertsson et al., 2014). We normalize the average firm size in the original steady state to 1 by setting the weight of labor in the households utility function $\bar{\phi}$ to 0.87. We set the exogenous destruction rate of firms δ

Table 5: Parameters

Parameter	Symbol	Value
Frisch elasticity	ϕ	0.5
Weight of labor	$\bar{\phi}$	0.87
Elasticity of intertemporal substitution	γ	1
Discount factor	β	0.99
Destruction rate	δ	0.0125
Pareto shape	k	6
Pareto scale	z_m	1
Elasticity of substitution (goods)	σ	5
Elasticity of substitution (employers)	η	16.5
Entry costs	f_e	0.99
Fixed costs	f_x	0.22

to 0.0125, based on evidence on the exit rates of establishments in Germany from Fackler et al. (2013). They report annual exit rates of roughly 0.05 for establishments older than five years. We use this number since we have no dimension of firm age in the model.[7] For lack of better evidence, we follow Colciago and Silvestrini (2022) by setting the entry costs f_e such that ξf_e is 1 in the original steady state and setting the fixed costs f_x such that $\frac{f_e}{f_x}$ equals 4.5. We further set the minimal productivity of the Pareto distribution z_m to 1 and its exponent k to 6. Both values have been recently used by Colciago and Silvestrini (2022) and Ascari et al. (2021). Finally, we set the elasticity of substitution for goods σ to 5 and the elasticity of substitution for employers η to 16.5. These values match the size wage premium for West Germany around the year 1990

[7]This implicitly means that in the model, a firm directly behaves as if it is an old firm after its entry.

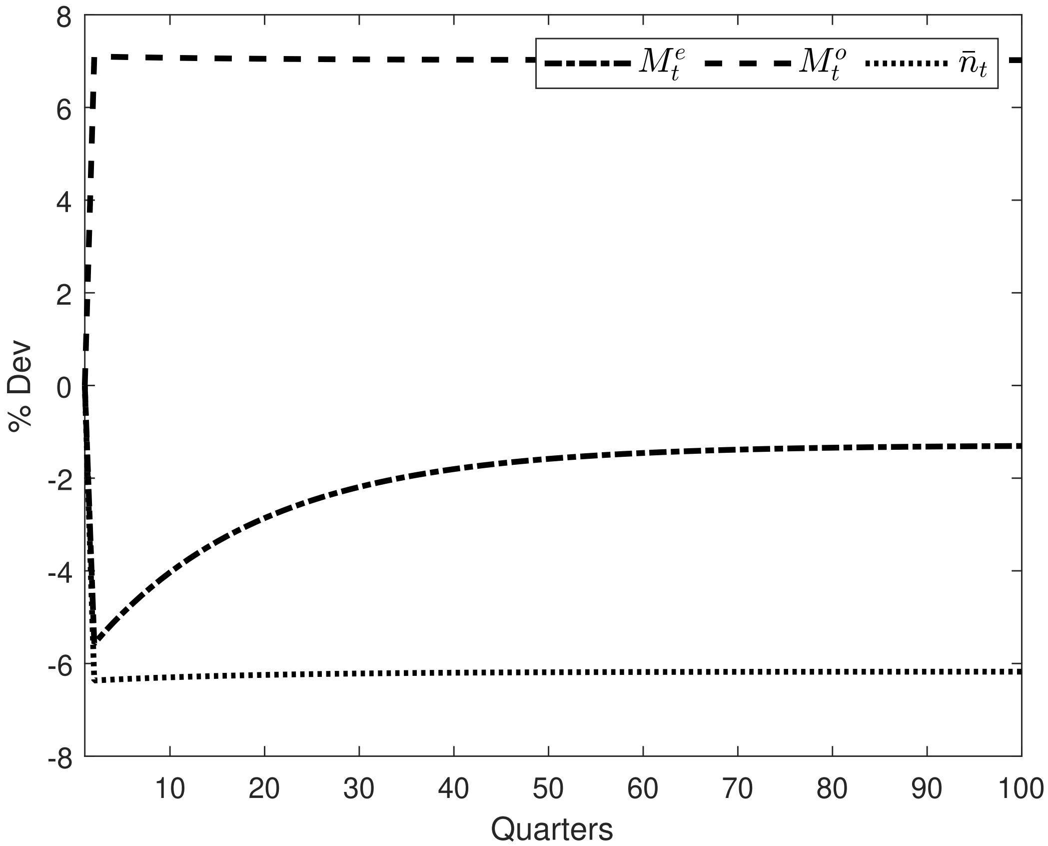
from the average wage regression (see Figure 4) and the overall markup in Germany from Christopoulou and Vermeulen (2012). Aggregate productivity Z is set to 1. Table 5 summarizes the model parameter values.

7.6 Model results

We simulate the empirically observed increase in the size wage premium by decreasing η (recall, $\frac{1}{\eta}$ corresponds to the size wage premium parameter) in period 1 and study the dynamic responses of the variables of interest. In the initial steady state (corresponding to the end of the 1980s), we set η to 16.5. This parameter choice corresponds to the observed size wage premium of 0.0606 in the 1980s (based on the average wage estimations). We decrease η to 13.5 to simulate the observed increase in the size wage premium to 0.0741.

In Figure 7, we report the model responses to this change in the size wage premium. The model economy responds with a decrease in the average establishment size $\bar{n}_t$ by about 6%. This magnitude corresponds very well to our fixed effects estimation results that suggest a decrease of average establishments size by 3.4% to 6.4% due to the increase in the size wage premium. We also see an increase in the number of operating firms M_t^o by about 7%. The increase in monopsony power increases the wage markdown the firms pay to the employees. As a result, a larger share of the existing firms decides to become active since they can earn positive profits. On the other hand, a larger number of active firms increases the relative price of the corporate good. This effective increase in fixed production costs dampens the increase in M_t^o. Similarly, creating a new firm effectively becomes more costly, which decreases M_t^e. It is important to emphasize that this observation in itself is no contradiction with our empirical observation of a higher number of entries, as this only means that there is less inflow into the overall mass of firms. The model equivalent to the empirically observed entries is the inflow of operating firms. Their model increase of 7% is somewhat larger than the empirically estimated increase of 1.2% to 5.2% due to the increase in the size wage premium between 1990 and the mid-2000s.

Figure 7: Model responses to an increase in the size wage premium

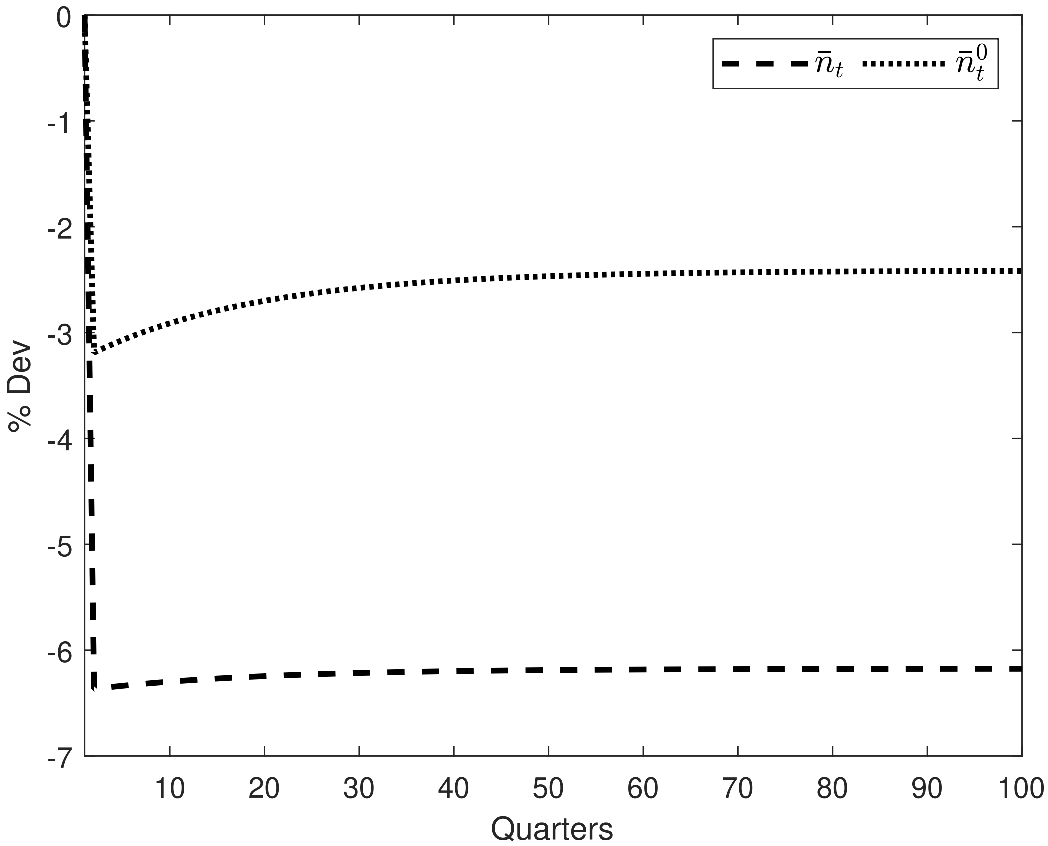


Notes: Model simulation of a decrease in η from 16.5 to 13.5. The variables are defined as follows: n_t is the average firm size, M_t^o is the number of operating firms and M_t^e is the number of new firms entering the market.

However, given that we do not target this outcome in the model, this result matches the empirical findings quite well.

Two channels drive the decline in the average firm size. We demonstrate these in Figure 8. Firstly, the steeper size-wage curve induces the operating firms to downsize by reducing employment. In Figure 8, we present this with the responses labeled by $\bar{n}_t^0$. Here we hold the threshold to become active z_0^c fixed, such that we only consider firms that would also be active in the previous steady state equilibrium. Hence, the response of $\bar{n}_t^0$ indicates the change in average firm size if the cut-off productivity had not changed. This response implies a decrease in the average establishment size by about 3% due to the downsizing of operating firms. The reason

Figure 8: The effect of the change in the productivity threshold on the average firm size

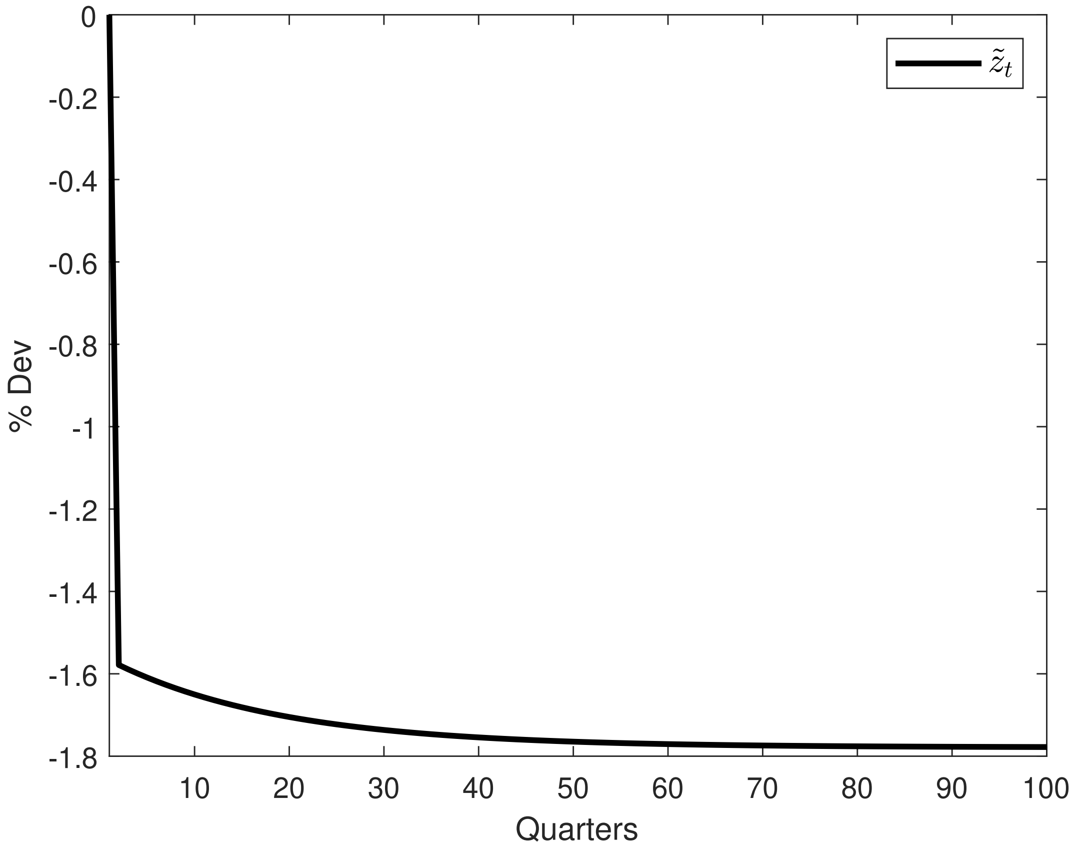


Notes: Model simulation of a decrease in η from 16.5 to 13.5. The variables are defined as follows: $\bar{n}_t$ is the average firm size, $\bar{n}_t^0$ is the average firm size holding the threshold to become active z_0^c fixed.

for their downsizing is that being large is more costly now. From the monopsony perspective, a steeper labor supply curve to the firm also implies a lower optimal employment level.

Secondly, since the actual productivity threshold of being active declines, less productive firms start to produce. These unproductive firms are smaller than incumbents and, hence, decrease the average firm size. As can be seen in Figure 8, about 50% of the overall decline in the average firm size is due to the downsizing of operating firms and 50% due to the entry of smaller firms. Idle firms becoming active due to the lower threshold are observationally equivalent to entries in the empirical part of the paper. Hence, the model results are consistent with the decomposition of the observed decline in the average establishment size, shown in Table 1.

Figure 9: Response of productivity to an increase in the size wage premium



Notes: Model simulation of a decrease in η from 16.5 to 13.5. n_t is the average firm size.

In Figure 9, we demonstrate the response of average productivity $\tilde{z}_t$ to the increase in the size wage premium in the model. We see a distinct decline in the average productivity in response to the increase in the size wage premium. Similar to the response of the average establishment size, there are two channels explaining this decline: First, the direct effect of unproductive firms entering the market and, second, the declining weight of productive incumbent firms. This model result is consistent with our empirical finding of an association between aggregate productivity and average establishment size.

All in all, our model can replicate the key empirical findings. We can theoretically demonstrate a negative link between average firm size and the size wage premium. Further, we show that this association is driven by the internalization of the size wage relation by operating firms and the entry of new firms. Moreover, our model predicts that the size wage premium is related to average productivity through average firm size. More precisely, average firm size is positively associated with average productivity. Hence, a decline in average firm size due to an increase in the size wage premium is associated with a decline in average productivity.

8 Conclusion

In this paper, we analyze the causes and consequences of the fall and rebound of the average establishment size in West Germany over the past 30 years. To this end, we first establish empirical associations between the average establishment size and aggregate productivity and between the average establishment size and the size wage premium by exploiting regional variation across West German counties. Secondly, we propose a dynamic model with heterogeneous firms and monopsonistic competition in the labor market, which is able to replicate our empirical findings. Moreover, we can use the model to investigate the economic mechanisms behind the empirical relationship we uncover. The key findings of the paper are the following:

1. We document a decline in the average establishment size in West Germany in the 1990s and early 2000s by close to two employees per establishment. This decline is followed by a strong rebound of the average establishment size by roughly 2.5 employees per establishment until 2019. In a decomposition exercise, we identify two sources driving these changes. Firstly, incumbent establishments downsized in the 1990s and early 2000s, while they grew in the years of the rebound. Secondly, an influx of new establishments reduced the average establishment size in the 1990s and early 2000s as these establishments tend to be considerably smaller than the average establishment. From the mid-2000s onwards, the number of new establishments declined markedly.

2. There is a positive and statistically significant association between aggregate productivity and the average establishment size. Our regional analysis reveals that the decrease in the average establishment size in the 1990s and early 2000s was associated with a decrease of GDP per capita by 2,000 to 2,500 euros, while the rebound was associated with an increase of GDP per capita by 2,900 to 3,800 euros.

3. We confirm the result of Colonnelli et al. (2018) and Lochner et al. (2020) that the size wage premium increased in West Germany in the 1990s and early 2000s and decreased afterwards. Furthermore, we find a negative and statistically significant association between average establishment size and the size wage premium on the regional level. Our results also show an economically meaningful and statistically significant association between the number of newly entering establishments per employee in a region and the regional size wage premium. These results imply that the fall and rebound in the average establishment size is associated with the increase and subsequent decrease in the size wage premium.

4. Building on these insights, we propose a dynamic general equilibrium model that is able to replicate the empirical results. The size wage premium reflects an upward-sloping labor supply curve for each firm. Therefore, our model is characterized by monopsony power in the labor market. Furthermore, the model features heterogeneous firms, endogenous entry of new firms and monopolistic competition on the goods market. Studying the model's responses to an increase in the size wage premium, we can infer the mechanics of the fall and rebound of the average establishment size. Incumbent establishments reduce their size as being large gets more costly. Equivalently, increasing monopsony power of the employers decreases their size as monopsonists reduce their employment to reduce their marginal labor costs (similar to the monopolist's reduction in output to maximize profits). Furthermore, the prospect of increasing monopsony profits draws in more unproductive and, therefore, small firms, which would be too unproductive to enter

a more competitive market. The increased number of unproductive firms and the lower weight of the productive firms in the economy decreases aggregate productivity.

These results imply that monopsony power incentivizes incumbent establishments to downsize and new establishments to enter, which reduces both the average establishment size and aggregate productivity. Put differently, the size wage relation is a source of misallocation, reallocating employees from large productive to small unproductive establishments. The reason is that it acts as an implicit tax on the size of establishments. Therefore, shutting down the size wage relation would increase establishment size and aggregate productivity. For instance, a minimum wage potentially cuts off a part of the size wage curve and prevents some misallocation from productive to unproductive firms. Other conceivable measures could be increasing workers' mobility, e.g., with the help of commuting allowances or more stringent working-from-home policies.

A Appendix

A.1 Average establishment size by economic sector

To get a more precise picture on the driving forces of the fall and rebound of the average establishment size, we split up the average size series by economic sectors. In the upper two panels of Figure A.1, we present the average establishment size in the manufacturing and the service sector. Clearly, the evolution of average establishment size shown in Figure 1 is driven by the manufacturing sector as the shapes of these two series are very similar. However, the average size of establishments in the service sector also declined in the 1990s, albeit less markedly and persistently than in manufacturing. Furthermore, the rise in the average establishment size in the service sectors started earlier than the rise in the manufacturing sector.

Turning to the average establishments size in large establishments, presented in the lower panels of Figure A.1, reveals additional insights. Again, the series for the manufacturing sector mirrors the corresponding series for all sectors in Figure 1. Interestingly, large establishment in the service sector reduced their size already in the 1980s and continued to do so until the mid 2000s. Hence, while there are marked differences in the evolution of the average establishment size between these two sectors, there is no indication that the fall and rebound is simply due to a changing sector composition.

Figure A.1: Average establishment size in manufacturing and services

(a) Average establishment size in manufacturing

(b) Average establishment size in services

(c) Average establishment size in large establishments in manufacturing

(d) Average establishment size in large establishments in services

A.2 Local labor market regions

In Table A.1, we present the results for the association of labor productivity and average establishment size employing local labor market according to Breidenbach et al. (2018) instead of counties. The main implications remain the unchanged. This is also true when we use a broader local labor market definition according to Kosfeld and Werner (2012), shown in Table A.2

Table A.1: Labor productivity and the average establishment size in narrower local labor markets

Dep. Var.:	GDP per capita			GDP per hour		
	OLS	FE	FE	OLS	FE	FE
Avg. size	894.13***	889.83***	905.62***	0.56***	0.54***	0.53***
	(111.48)	(148.74)	(154.27)	(0.05)	(0.17)	(0.18)
Share rout.	-147.15***	-4.44	17.87	-0.03	0.01	0.03
	(41.29)	(16.35)	(23.30)	(0.03)	(0.02)	(0.02)
Share fem.	70.10	178.55**	204.86***	-0.10	0.13	0.18**
	(126.49)	(90.24)	(78.51)	(0.10)	(0.09)	(0.08)
Share for.	326.30***	244.97**	65.46	0.30***	0.18*	-0.09
	(119.96)	(100.75)	(100.92)	(0.10)	(0.10)	(0.10)
Share hq.	434.27***	3.72	98.87	0.61***	-0.02	0.05
	(139.68)	(106.35)	(105.81)	(0.11)	(0.11)	(0.11)
Share lq.	-346.17***	-293.63***	-18.97	-0.25**	-0.27***	-0.05
	(131.02)	(64.61)	(74.52)	(0.12)	(0.08)	(0.09)
Time Dummies	✓	✓	✓	✓	✓	✓
State Dummies	✓	-	-	✓	-	-
State-Time Int.	-	-	✓	-	-	✓
Sector Shares	-	-	✓	-	-	✓
Constant	27119.17***	15563.50***	10855.65*	40.43***	35.13***	26.98***
	(5562.34)	(3773.98)	(6342.45)	(4.42)	(3.92)	(6.83)
R^2	0.70	0.70	0.75	0.66	0.57	0.64
N	5058	5058	5058	4040	4040	4040

Notes: The observational unit is a local labor market according to Breidenbach et al. (2018). Robust standard errors clustered on the local labor market level are displayed in parentheses. *** p<0.01, ** p<0.05, * p<0.1 indicate the significance levels. Avg. size: average establishment size, share rout.: share of routine workers, share fem.: share of female workers, share for.: share of foreign workers, share hq.: share of high qualified workers, share lq.: share of low qualified workers, and state-time int.: state-time interactions.

Table A.2: Labor productivity and the average establishment size in broader local labor markets

Dep. Var.:	GDP per capita			GDP per hour		
	OLS	FE	FE	OLS	FE	FE
Avg. size	669.80***	1097.44***	1112.28***	0.53***	0.85***	0.82***
	(133.02)	(148.59)	(158.63)	(0.10)	(0.17)	(0.18)
Share rout.	-213.00***	-30.56*	12.65	-0.05	0.01	0.05
	(59.94)	(17.23)	(32.47)	(0.05)	(0.02)	(0.03)
Share fem.	-17.36	65.26	65.19	-0.06	0.07	0.12
	(153.36)	(149.93)	(134.32)	(0.12)	(0.13)	(0.13)
Share for.	377.15**	346.37**	212.64	0.38***	0.21	0.00
	(155.24)	(174.25)	(200.02)	(0.14)	(0.15)	(0.19)
Share hq.	335.38*	82.15	180.89	0.59***	-0.01	0.08
	(201.05)	(133.49)	(154.45)	(0.17)	(0.16)	(0.18)
Share lq.	-269.06**	-335.84***	-101.01	-0.25*	-0.34***	-0.18
	(122.52)	(88.81)	(96.59)	(0.14)	(0.10)	(0.15)
Time Dummies	✓	✓	✓	✓	✓	✓
State Dummies	✓	-	-	✓	-	-
State-Time Int.	-	-	✓	-	-	✓
Sector Shares	-	-	✓	-	-	✓
Constant	36253.61***	17873.99***	9454.59	40.38***	34.58***	36.41***
	(6093.98)	(4195.75)	(7533.00)	(5.37)	(4.36)	(8.46)
R^2	0.73	0.75	0.82	0.72	0.66	0.71
N	2710	2710	2710	2160	2160	2160

Notes: The observational unit is a local labor market according to Kosfeld and Werner (2012). Robust standard errors clustered on the local labor market level are displayed in parentheses. *** p<0.01, ** p<0.05, * p<0.1 indicate the significance levels. Avg. size: average establishment size, share rout.: share of routine workers, share fem.: share of female workers, share for.: share of foreign workers, share hq.: share of high qualified workers, share lq.: share of low qualified workers, and state-time int.: state-time interactions.

The analysis of the association of the average establishment size and the size wage premium is robust to different local labor market definitions, too. The results with the definitions according to Breidenbach et al. (2018) and Kosfeld and Werner (2012) are displayed in Table A.3 and Table A.3, respectively. However, when we use the broader local labor market definition of Kosfeld and Werner (2012), the number of observations reduces notably. Therefore, we lose statistical power and, hence, we see statistically insignificant coefficients in some cases, even though the point estimate of these parameters is in the range of the main estimates.

Table A.3: Average establishment size and the size wage premium in narrower local labor markets

Dependent variable: average establishment size

	AKM based			average wage based		
	OLS	FE	FE	OLS	FE	FE
swp_{rt}	-142.68	-87.88***	-74.07**	-71.05	-61.57***	-57.52***
	(113.49)	(32.90)	(30.86)	(66.76)	(14.23)	(12.36)
Share rout.	55.70	-29.26**	-34.49**	24.83	-7.78	-12.74*
	(39.56)	(13.12)	(14.31)	(30.23)	(5.86)	(7.62)
Share fem.	-279.14***	-43.87**	-19.18	-296.30***	-84.82***	-50.89***
	(67.06)	(17.56)	(19.26)	(69.62)	(18.36)	(18.18)
Share for.	41.06	95.91**	104.77***	24.32	67.56***	69.32***
	(56.27)	(40.51)	(37.57)	(39.84)	(14.23)	(16.28)
Share hq.	300.08***	6.73	7.04	245.23***	19.71	7.28
	(87.99)	(29.78)	(32.93)	(71.60)	(20.82)	(21.18)
Share lq.	47.99	-0.37	-1.17	74.09*	17.31	6.43
	(45.69)	(20.40)	(21.78)	(42.00)	(18.15)	(18.73)
Time Dummies	✓	✓	✓	✓	✓	✓
State Dummies	✓	-	-	✓	-	-
State-Time Int.	-	-	✓	-	-	✓
Sector Shares	-	-	✓	-	-	✓
Constant	73.31***	88.12***	87.57***	89.88***	77.16***	85.04***
	(14.41)	(9.91)	(15.97)	(14.09)	(6.00)	(15.16)
R^2	0.33	0.26	0.34	0.32	0.18	0.39
N	1010	1010	1010	7070	7070	7070

Notes: The observational unit is a local labor market according to Breidenbach et al. (2018). Robust standard errors clustered on the local labor market level are displayed in parentheses. *** p<0.01, ** p<0.05, * p<0.1 indicate the significance levels. swp_{rt}: size wage premium, share rout.: share of routine workers, share fem.: share of female workers, share for.: share of foreign workers, share hq.: share of high qualified workers, share lq.: share of low qualified workers, and state-time int.: state-time interactions.

Table A.4: Average establishment size and the size wage premium in broader local labor markets
Dependent variable: average establishment size

	AKM based			average wage based		
	OLS	FE	FE	OLS	FE	FE
swp_{rt}	-75.11	-121.14**	-80.50	-187.67**	-39.44**	-29.74*
	(73.79)	(57.18)	(49.31)	(86.51)	(16.58)	(16.45)
Share rout.	66.88	-20.17	-32.60	23.38	-3.21	-7.11
	(52.71)	(18.47)	(23.02)	(34.70)	(8.04)	(10.88)
Share fem.	-224.76***	-34.34*	6.12	-236.79***	-83.79***	-33.91*
	(65.98)	(20.17)	(18.85)	(70.58)	(20.50)	(19.44)
Share for.	52.79	108.01*	107.92*	12.88	64.70***	58.33***
	(46.38)	(57.58)	(55.97)	(32.47)	(20.92)	(19.62)
Share hq.	305.65**	25.79	29.20	234.18**	2.56	-0.98
	(129.16)	(40.54)	(47.35)	(102.25)	(15.48)	(17.46)
Share lq.	35.72	-4.69	-1.14	64.52	29.80	22.35
	(46.21)	(25.73)	(28.02)	(40.87)	(19.95)	(21.42)
Time Dummies	✓	✓	✓	✓	✓	✓
State Dummies	✓	-	-	✓	-	-
State-Time Int.	-	-	✓	-	-	✓
Sector Shares	-	-	✓	-	-	✓
Constant	50.34**	78.73***	71.36***	84.53***	70.36***	63.89***
	(19.52)	(12.84)	(19.56)	(13.45)	(6.56)	(14.23)
R^2	0.34	0.34	0.44	0.35	0.26	0.49
N	540	540	540	3780	3780	3780

Notes: The observational unit is a local labor market according to Kosfeld and Werner (2012). Robust standard errors clustered on the local labor market level are displayed in parentheses. *** p<0.01, ** p<0.05, * p<0.1 indicate the significance levels. swp_{rt}: size wage premium, share rout.: share of routine workers, share fem.: share of female workers, share for.: share of foreign workers, share hq.: share of high qualified workers, share lq.: share of low qualified workers, and state-time int.: state-time interactions.

Furthermore, also the results concering the relationship of establishments entries and the size wage premium are robust if we employ the two local labor market definitions, see Tables A.5 and A.4. Again, some estimates based on the local labor market definition of Kosfeld and Werner (2012) are statistically insignificant due to the lower number of observations.

Table A.5: Entries per employee and the size wage premium in narrower local labor markets
Dependent variable: entries per employee

	AKM based			average wage based		
	OLS	FE	FE	OLS	FE	FE
swp_{rt}	0.181**	0.135***	0.115***	0.033***	0.016***	0.014***
	(0.078)	(0.039)	(0.037)	(0.006)	(0.003)	(0.003)
Share rout.	-0.023	-0.002	0.016	-0.000	-0.002*	-0.001
	(0.028)	(0.013)	(0.015)	(0.002)	(0.001)	(0.002)
Share fem.	0.259***	0.068***	0.058**	0.034***	0.015***	0.012***
	(0.038)	(0.021)	(0.023)	(0.005)	(0.003)	(0.003)
Share for.	0.113*	-0.127***	-0.136***	0.016***	-0.007**	-0.012***
	(0.060)	(0.027)	(0.029)	(0.006)	(0.003)	(0.003)
Share hq.	-0.249***	0.017	0.015	-0.023***	0.000	-0.002
	(0.061)	(0.028)	(0.033)	(0.005)	(0.003)	(0.004)
Share lq.	-0.214***	-0.073**	-0.051*	-0.030***	-0.013***	-0.008**
	(0.050)	(0.028)	(0.027)	(0.006)	(0.004)	(0.003)
Time Dummies	✓	✓	✓	✓	✓	✓
State Dummies	✓	-	-	✓	-	-
State-Time Int.	-	-	✓	-	-	✓
Sector Shares	-	-	✓	-	-	✓
Constant	0.089***	0.077***	0.038	0.007***	0.008***	0.003
	(0.023)	(0.013)	(0.028)	(0.002)	(0.001)	(0.003)
R^2	0.47	0.59	0.68	0.54	0.64	0.70
N	1010	1010	1010	7070	7070	7070

Notes: The observational unit is a local labor market according to Breidenbach et al. (2018). Robust standard errors clustered on the county level are displayed in parentheses. *** p<0.01, ** p<0.05, * p<0.1 indicate the significance levels. swp_{rt}: size wage premium, share rout.: share of routine workers, share fem.: share of female workers, share for.: share of foreign workers, share hq.: share of high qualified workers, share lq.: share of low qualified workers, and state-time int.: state-time interactions.

Table A.6: Entries per employee and the size wage premium in broader local labor markets
Dependent variable: entries per employee

	AKM based			average wage based		
	OLS	FE	FE	OLS	FE	FE
swp_{rt}	0.229*	0.150**	0.082	0.050***	0.021***	0.016***
	(0.130)	(0.069)	(0.064)	(0.008)	(0.006)	(0.004)
Share rout.	-0.047	-0.001	0.011	-0.001	-0.002	-0.002
	(0.035)	(0.018)	(0.020)	(0.002)	(0.001)	(0.002)
Share fem.	0.211***	0.067*	0.062*	0.028***	0.015***	0.014***
	(0.058)	(0.035)	(0.037)	(0.007)	(0.005)	(0.004)
Share for.	0.046	-0.119***	-0.140***	0.011*	-0.003	-0.008**
	(0.066)	(0.032)	(0.036)	(0.007)	(0.004)	(0.003)
Share hq.	-0.220***	-0.004	0.007	-0.019***	0.002	-0.000
	(0.071)	(0.033)	(0.038)	(0.005)	(0.003)	(0.004)
Share lq.	-0.182***	-0.042	-0.015	-0.026***	-0.011***	-0.007*
	(0.056)	(0.027)	(0.035)	(0.006)	(0.003)	(0.004)
Time Dummies	✓	✓	✓	✓	✓	✓
State Dummies	✓	-	-	✓	-	-
State-Time Int.	-	-	✓	-	-	✓
Sector Shares	-	-	✓	-	-	✓
Constant	0.100***	0.064***	0.050	0.006**	0.007***	0.002
	(0.028)	(0.016)	(0.031)	(0.003)	(0.001)	(0.003)
R^2	0.48	0.65	0.73	0.59	0.72	0.78
N	540	540	540	3780	3780	3780

Notes: The observational unit is a local labor market according to Kosfeld and Werner (2012). Robust standard errors clustered on the county level are displayed in parentheses. *** p<0.01, ** p<0.05, * p<0.1 indicate the significance levels. swp_{rt}: size wage premium, share rout.: share of routine workers, share fem.: share of female workers, share for.: share of foreign workers, share hq.: share of high qualified workers, share lq.: share of low qualified workers, and state-time int.: state-time interactions.

A.3 Derivation of the decomposition of a change in the average establishment size

Consider the the definition of the average establishment size

$$\bar{n}_t = \frac{1}{M_t} \sum_{i=1}^{M_t} n_{i,t}$$

where M_t is the number of establishment in t. Consider three types of establishments: incumbents C (exist in t and $t+1$), entries E (exist only in $t+1$), and exits X (exits only in t). Define the average establishment size of a specific type $j \in \{C, X, E\}$ as

$$\bar{n}_t^j = \frac{1}{M_t^j} \sum_{i=1}^{M_t^j} n_{i,t}^j$$

and the share of an establishment type in all establishments as

$$\theta_t^j = \frac{M_t^j}{M_t}.$$

Using these definitions, we can write the average establishment size in t as

$$\bar{n}_t = \theta_t^C \bar{n}_t^C + \theta_t^X \bar{n}_t^X$$

and the average establishment size in $t+1$ as

$$\bar{n}_{t+1} = \theta_{t+1}^C \bar{n}_{t+1}^C + \theta_{t+1}^E \bar{n}_{t+1}^E.$$

Thus, an absolute change in the average establishment size can be expressed as

$$\bar{n}_{t+1} - \bar{n}_t = \theta_{t+1}^C \bar{n}_{t+1}^C - \theta_t^C \bar{n}_t^C + \theta_{t+1}^E \bar{n}_{t+1}^E - \theta_t^X \bar{n}_t^X.$$

Adding and subtracting $\theta_t^C n_{t+1}^C$ on the left-hand side yield

$$\begin{aligned}
\Delta\bar{n}_{t+1} &= \theta_{t+1}^C \bar{n}_{t+1}^C - \theta_t^C \bar{n}_t^C + \theta_{t+1}^E \bar{n}_{t+1}^E - \theta_t^X \bar{n}_t^X + \theta_t^C n_{t+1}^C - \theta_t^C n_{t+1}^C \\
\Delta\bar{n}_{t+1} &= \theta_t^C [\bar{n}_{t+1}^C - \bar{n}_t^C] + [\theta_{t+1}^C - \theta_t^C] \bar{n}_t^C + [\theta_{t+1}^C - \theta_t^C][\bar{n}_{t+1}^C - \bar{n}_t^C] + \theta_{t+1}^E \bar{n}_{t+1}^E - \theta_t^X \bar{n}_t^X
\end{aligned}$$

Adding and subtracting $\bar{n}$ on the left-hand side yield

$$\Delta\bar{n}_{t+1} = \theta_t^C[\bar{n}_{t+1}^C - \bar{n}_t^C] + [\theta_{t+1}^C - \theta_t^C]\bar{n}_t^C + [\theta_{t+1}^C - \theta_t^C][\bar{n}_{t+1}^C - \bar{n}_t^C] + \theta_{t+1}^E\bar{n}_{t+1}^E - \theta_t^X\bar{n}_t^X + \bar{n}_t - \bar{n}_t$$

$$\begin{aligned}\Delta\bar{n}_{t+1} &= \theta_t^C[\bar{n}_{t+1}^C - \bar{n}_t^C] + [\theta_{t+1}^C - \theta_t^C]\bar{n}_t^C + [\theta_{t+1}^C - \theta_t^C][\bar{n}_{t+1}^C - \bar{n}_t^C] \\ &\quad + \theta_{t+1}^E\bar{n}_{t+1}^E - \theta_t^X\bar{n}_t^X + \frac{M_t^C + M_t^X}{M_t}\bar{n}_t - \frac{M_{t+1}^C + M_{t+1}^E}{M_{t+1}}\bar{n}_t \\ \Delta\bar{n}_{t+1} &= \theta_t^C[\bar{n}_{t+1}^C - \bar{n}_t^C] + [\theta_{t+1}^C - \theta_t^C]\bar{n}_t^C + [\theta_{t+1}^C - \theta_t^C][\bar{n}_{t+1}^C - \bar{n}_t^C] \\ &\quad + \theta_{t+1}^E\bar{n}_{t+1}^E - \theta_t^X\bar{n}_t^X + \theta_t^C\bar{n}_t + \theta_t^X\bar{n}_t - \theta_{t+1}^C\bar{n}_t - \theta_{t+1}^E\bar{n}_t.\end{aligned}$$

Collecting terms yields the final equation of the decomposition

$$\Delta\bar{n}_{t+1} = \theta_t^C\Delta\bar{n}_{t+1}^C + \Delta\theta_{t+1}^C(\bar{n}_t^C - \bar{n}_t) + \Delta\theta_{t+1}^C\Delta\bar{n}_{t+1}^C + \theta_{t+1}^E(\bar{n}_{t+1}^E - \bar{n}_t) - \theta_t^X(\bar{n}_t^X - \bar{n}_t).$$

A.4 Model derivations

A.4.1 Household

As described in the main text, the household solves the maximization problem

$$\max_{C_t, N_t, b_{t+1}, x_{t+1}} E_0 \sum_{t=0}^{\infty} \beta^t \left[\frac{C_t}{1-\gamma}^{1-\gamma} - \bar{\varphi}^{\frac{1}{\varphi}} \frac{N_t}{1+\frac{1}{\varphi}}^{1+\frac{1}{\varphi}} \right]$$

$$s.t.$$

$$b_{t+1} + \tilde{v}_t(M_t + M_t^E)x_{t+1} + C_t = (1 + r_t)b_t + \tilde{v}_t M_t x_t + \tilde{d}_t M_t^o x_t + W_t N_t.$$

The FOCs are:

$$
\begin{aligned}
&\partial C_t: \; C_t^{-\gamma} - \lambda_t = 0 \\
&\Leftrightarrow C_t^{-\gamma} = \lambda_t \\
&\partial N_t: \; -\bar{\varphi}^{\frac{1}{\varphi}} N_t^{\frac{1}{\varphi}} + \lambda_t W_t = 0 \\
&\Leftrightarrow W_t = \bar{\varphi}^{\frac{1}{\varphi}} \frac{N_t^{\frac{1}{\varphi}}}{C_t^{-\gamma}} \\
&\partial b_{t+1}: -\lambda_t + E_t \beta (1 + r_{t+1}) \lambda_{t+1} = 0 \\
&\Leftrightarrow C_t^{-\gamma} = E_t \beta (1 + r_{t+1}) C_{t+1}^{-\gamma} \\
&\partial x_{t+1}: -\tilde{v}_t (M_t + M_t^E) \lambda_t + E_t \beta (\tilde{v}_{t+1} M_{t+1} + \tilde{d}_{t+1} M_{t+1}^o) \lambda_{t+1} = 0 \\
&\Leftrightarrow \tilde{v}_t = E_t \beta (1 - \delta)((1 - G(z_{t+1}^c)) \tilde{d}_{t+1} + \tilde{v}_{t+1}) \left(\frac{C_{t+1}}{C_t} \right)^{-\gamma},
\end{aligned}
$$

where the law of motion for the number of firms was used in the last step:

$$M_t = (1 - \delta)(M_{t-1} + M_{t-1}^E),$$

together with the definition of the number of operating firms:

$$M_t^o = (1 - G(z_t^c)) M_t.$$

A.4.2 Firms

To derive the pricing rule of a firm we first write real profits

$$\pi_t(z) = q_t(z) \rho_t(z) - w_t(z) n_t(z) - \xi_t f_x$$

as a function of only the firms output $q_t(z)$ and its price $\rho_t(z)$ using equations (15) and (26). Now the maximization problem is

$$\max_{q_t(z)} \pi_t(z) = q_t(z)\rho_t(z) - \left(\frac{q_t(z)}{Z_t z}\right)^{\frac{\eta+1}{\eta}} N_t^{-\frac{1}{\eta}} W_t - \xi_t f_x$$

$$\partial q_t(z) : \rho_t(z) + q_t(z)\frac{\partial \rho_t(z)}{\partial q_t(z)} - \frac{\eta+1}{\eta}\left(\frac{q_t(z)}{Z_t z}\right)^{\frac{1}{\eta}} \frac{1}{Z_t z} N_t^{-\frac{1}{\eta}} W_t = 0.$$

Reinserting equations (15) and (26) gives us the pricing rule (28)

$$\rho_t(z) = \frac{\sigma}{\sigma-1}\frac{\eta+1}{\eta}\frac{w_t(z)}{Z_t z}.$$

Plugging (28) and (23) into (15) gives us the wage of a firm with individual productivity z

$$w_t(z) = \left(\frac{\sigma}{\sigma-1}\frac{\eta+1}{\eta}\right)^{-\frac{\sigma}{\eta+\sigma}} (Z_t z)^{\frac{\sigma-1}{\eta+\sigma}} \left(\frac{D_t}{N_t}\right)^{\frac{1}{\eta+\sigma}} W_t^{\frac{\eta}{\eta+\sigma}}.$$

Plugging this into

$$\pi_t^o(z) = \rho_t(z) q_t(z) - w_t(z) n_t(z)$$

together with (28) and (23) and using straight forward algebra gives us

$$\pi_t^o(z) = \left(\frac{\sigma}{\sigma-1}\frac{\eta+1}{\eta}\right)^{1-\sigma} \left[1 - \frac{\sigma-1}{\sigma}\frac{\eta}{\eta+1}\right] \left(\frac{w_t(z)}{Z_t z}\right)^{1-\sigma} D_t$$

$$\pi_t^o(z) = \frac{\sigma+\eta}{\sigma(\eta+1)}\left(\frac{\sigma}{\sigma-1}\frac{\eta+1}{\eta}\right)^{1-\sigma} D_t (Z_t z)^{\sigma-1} \left[\left(\frac{\sigma}{\sigma-1}\frac{\eta+1}{\eta}\right)^{-\frac{\sigma}{\eta+\sigma}} (Z_t z)^{\frac{\sigma-1}{\eta+\sigma}} \left(\frac{D_t}{N_t}\right)^{\frac{1}{\eta+\sigma}} W_t^{\frac{\eta}{\eta+\sigma}}\right]^{1-\sigma}$$

$$\pi_t^o(z) = \left[\frac{((\sigma-1)\eta)^{\eta(\sigma-1)} D_t^{\eta+1} \left(\frac{N_t}{W_t^{\eta}}\right)^{\sigma-1}}{(\sigma(\eta+1))^{\sigma(\eta+1)}}\right]^{\frac{1}{\sigma+\eta}} (\sigma+\eta)(Z_t z)^{\frac{(\eta+1)(\sigma-1)}{\eta+\sigma}}.$$

A.4.3 Average Productivity and Equilibrium

The starting point for equation (39) is the aggregate price index for the consumption bundle

$$P_t = \left(\int_{z_t^c}^{\infty} p_t(z)^{1-\sigma} M_t^o \frac{g(z)}{(1 - G(z_t^c))} dz \right)^{\frac{1}{1-\sigma}}.$$

Dividing the equation by P_t and using (28) and (29) gives us

$$\begin{aligned} 1 = {M_t^o}^{\frac{1}{1-\sigma}} &\left(\frac{\sigma}{\sigma - 1} \frac{\eta + 1}{\eta} \right) \times \\ &\left[\left(\frac{\sigma}{\sigma - 1} \frac{\eta + 1}{\eta} \right)^{-\frac{\sigma}{\eta+\sigma}} \left(\frac{D_t}{N_t} \right)^{\frac{1}{\eta+\sigma}} W_t^{\frac{\eta}{\eta+\sigma}} \right] \times \\ &\quad Z_t^{-\frac{\eta+1}{\eta+\sigma}} \left(\left(\int_{z_t^c}^{\infty} z^{\kappa} \frac{g(z)}{(1 - G(z_t^c))} dz \right)^{\frac{1}{\kappa}} \right)^{-\frac{\eta+1}{\eta+\sigma}}. \end{aligned}$$

Using (36), (29) and (28) gives us (39).

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